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MEDICAL CARE
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Diverging Trends in Hospitalization: Fact or Artifact?

HAROLD S. LUFT, PH.D.*

It has been noticed recently that hospital use by Blue Cross enrollees has fallen substantially in the last decade and that this decline is in marked contrast to relatively stable utilization reported by hospitals. The trends in length of stay in the two data sources are almost identical—all the differential is attributable to differing trends in admission rates: a sharp increase for all persons under 65 years of age in contrast to a moderate decline for Blue Cross members. These two trends are diverging at the rate of 2 per cent per year. Household survey data for the under-65 population with private health insurance also indicates an increasing admission trend, in contrast to the Blue Cross figures, even though Blue Cross covers half the insured population. About half of the differential trend can be accounted for by the increasing rate of duplicate coverage, which artificially increases the enrollments reported by carriers. The increased duplicate coverage is a reflection of more multiple-earner families, a shifting occupation mix, better fringe benefits, higher unemployment and smaller families. Part of the increased admissions for the total under-65 population is due to Medicare coverage for the disabled and to Medicaid. Most important, it is apparent that existing data provide an inadequate and potentially misleading picture of the most crucial measures of utilization.

ATTENTION has recently been drawn to the substantial decline in hospital use over the last decade for members of both Health Maintenance Organizations (HMOs) and Blue Cross plans.[1–3] The number of inpatient days for Blue Cross regular group members, that is excluding those with nongroup or Medicare complementary coverage, has fallen from 865 days per thousand in 1968 to 708 in 1978, a decline of 18 per cent.† This decline is in contrast to a rising trend in hospitalization for the overall population. Figure 1 presents the basic data concerning the rate of hospital use for Blue Cross regular group members in contrast to the rates for (1) the total popu-

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† Two segments of the Blue Cross membership are excluded from these data: 1) persons with Medicare complementary coverage and 2) those with individually purchased policies. The regular group enrollment and utilization data exclude Washington State and Montana. Medicare complementary and individual enrollees are relatively high users of hospital care. More important, changes over time in their proportion of the Blue Cross enrollment would be reflected in the average utilization rate for all Blue Cross members.

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lation using nonfederal, short-term general and other special hospitals as reported by the American Hospital Association (AHA) and 2) the under-65 noninstitutionalized population using those hospitals, as reported by the National Hospital Discharge Survey (NHDS).[4–6]

There are important policy implications if this divergence is real, but this article will present evidence that much of the apparent reduction over time in hospital use by Blue Cross members is a statistical artifact. Moreover, the factors creating this artifact are likely to account for the similar downward trends in utilization among enrollees in other conventional insurance plans and in HMOs. Data artifacts may lead to incorrect analyses and subsequent policies. More important, some of the artifacts arise from real trends of substantial policy importance in themselves.

This study has certain aspects of a detective story with several culprits and no motives. The data in Figure 1 indicate clearly diverging trends in hospital use. The task at hand is to indentify plausible explanations for the divergence and examine the evidence to support, reject or partially accept each explanation. Unfortunately, most of the evidence is circumstantial and some readers are likely to be more convinced than others. Even I am not willing to consider the case closed and I welcome new evidence for alternative explanations.

The body of this article is divided into four sections. The first attempts to clarify those aspects of hospitalization that exhibit diverging trends. This involves decompos-

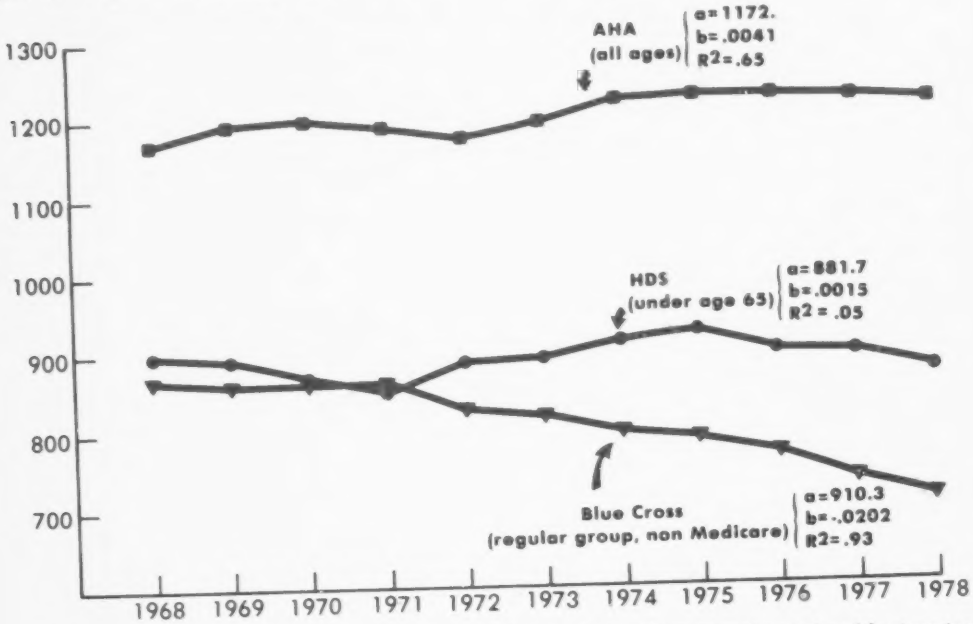

FIG. 1. Inpatient days per 1,000 persons. The AHA data are derived from annual statistics of the American Hospital Association and U.S. Census estimates of the total civilian population. National Hospital Discharge Survey (NHDS) data are from the NCHS survey of nonfederal short-stay hospitals, excluding newborns. Blue Cross data are for regular group members, excluding those with Medicare complementary coverage.

The standard formula for an annual rate of change is $y_t = a(1 + b)^t$, where y_t is the observed value in year t, a is a constant, b is the rate of change, and t is a time variable, equal to 1 in 1968, . . . 11 in 1978. The equation is linear in logs: $\log y_t = \log a + t \log (1 + b)$. The intercept, a, is the estimated value for 1967 (i.e., t = 0) and b represents the annual proportionate rate of change. For instance, b = −0.0202 means an annual rate of decrease of 2.02 per cent.

ing total hospital utilization rates into admission rates and length of stay, as well as excluding the over-65-population from the national data. Within this breakdown, it is important to determine whether hospitalization trends for Blue Cross enrollees are substantially different from the trends for people with commercial insurance coverage. The second section of the article examines the evidence of an artifact created by duplicate private insurance coverage that might explain part of the diverging trends in hospitalization. The third section introduces additional factors that may account for a portion of the remaining unexplained difference. A final section outlines some of the implications of the duplicate coverage problem for policy and analysis.

Clarifying the Sources of Divergence

The overall utilization data in Figure 1 can be split into trends in length of stay and admissions (or discharges) per 1,000.. As indicated in Figure 2, both BC and NHDS data show the average length of stay declining at about the same rate, 1.1–1.3 per cent per year. (This rate of change is derived from a regression line through the points for each year, and thus reduces the impact of year-to-year variability.)‡§ Figure 3, however, shows a very different pattern with a clear *increase* in discharges for the under-65 population as a whole (the NHDS data), in contrast to a more modest, but clear *reduction* in the Blue Cross data. The two trend lines for admissions are diverging by about 2 per cent per year, that is, the NHDS rate is increasing by 1.30 per cent per year while the BC rate is falling by

‡ See the notes to Figure 1.

§ Some of the year-to-year variability in the NHDS is the result of its design. Since new hospitals are continually opening but the NHDS sampling frame is updated only every few years, the further one is from an updating year (1965, 1972 or 1975), the more the NHDS underestimates the true discharge rate.[7]

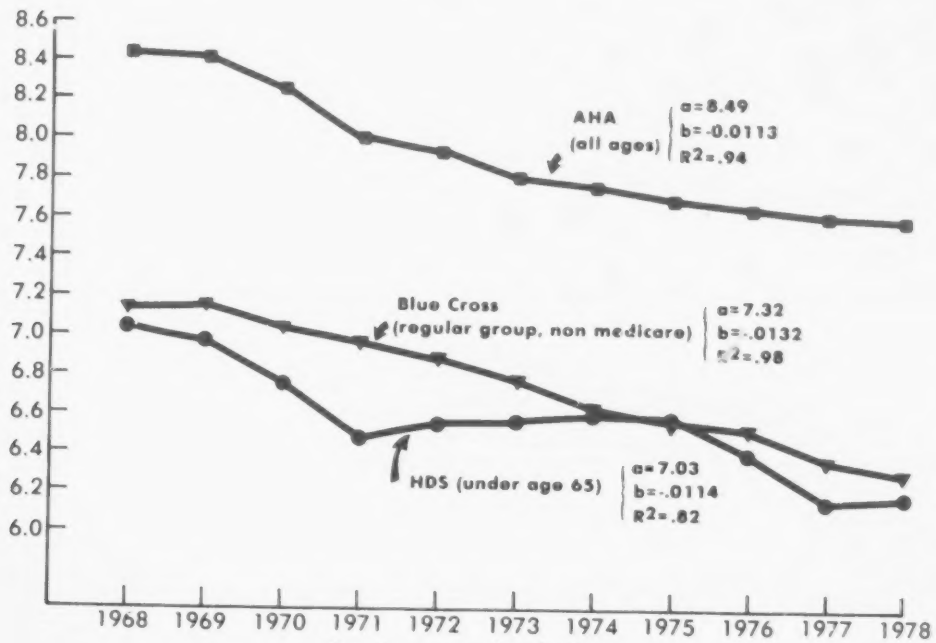

FIG. 2. Average length of stay.

0.67 per cent. The strong similarity in both the levels and trends, for length-of-stay supports the notion that the two data sets are capturing similar types of and changes in hospital use. Since the LOS calculation depends only on patient days and admissions, it is not subject to errors in the estimation of enrollments. This suggests that the impressive reduction in stay over the past decade is not only real, but has been experienced by both Blue Cross and non–Blue Cross members.[†] It is possible that Blue Cross enrollees experienced an even greater reduction in case-specific length of stay that was offset by changes in case mix. However, there is no evidence to support or reject this possibility.

An initial criticism of this analysis is that the NHDS includes all discharges from short-term non-Federal hospitals, regardless of coverage. By limiting the analysis to the population under age 65, the confounding influence of most Medicare beneficiaries can be eliminated, although under-65 Medicare disabled beneficiaries are included. Blue Cross accounts for about a third of all discharges among the nonelderly population. Therefore, if all the apparent BC decline were true, the average annual increase in admissions for the non-BC segment of the under-65 population would be 2.27 per cent, in contrast to the 1.30 per cent seen from the NHDS (Fig. 3).[#] If this increase is attributed to a subset of the non-BC population, the rate of divergence becomes even greater. For such a rapid increase in hospital admissions among the non-BC commercially insured

[†] For instance, during the period 1968 to 1977 the average length of stay for enrollees in the two California Kaiser plans fell at the rate of 1.56 per cent per year. This somewhat more rapid decline may be attributable to several factors, including the relatively greater importance of maternity admissions in HMO populations.

[#] This calculation is as follows: the one-third of the under-65 NHDS total represented by BC has an annual decline of 0.67 per cent while the total under-65 NHDS is increasing by 1.30 per cent. Therefore, letting x = annual change for the non-BC population,

$$0.33(-0.67) + 0.67x = 1.30$$
$$0.67x = 1.52$$
$$x = 2.27.$$

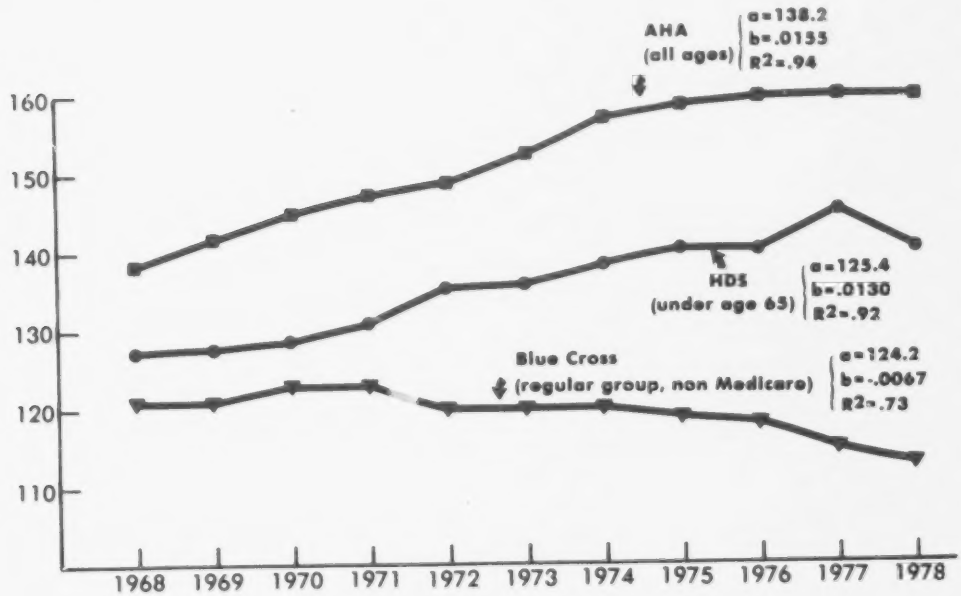

FIG. 3. Admissions or discharges per 1,000 persons.

population to have gone unnoticed and unreported would be surprising, but not impossible. The limited availability of data suggests that it may be difficult to monitor such changes in true hospital-use rates.

Although hospital utilization data for people with commerical coverage are not available, there is an indirect test of the possibility that a falling Blue Cross hospitalization rate is being offset by a rising commercially insured rate to produce the observed NHDS results. If this were true, it would probably be reflected in claims data. (The major reason to expect different patterns for Blue Cross and commercial enrollees is a "catching-up" in improved coverage among the latter.) In fact, for 1968-73, the period during which comparable data are available, hospital benefits per Blue Cross enrollee rose at an annual rate of 11.4 per cent in contrast to 8.2 per cent for commercially covered people.[8,9] To the extent that the charges paid by commercial carriers rose more rapidly than costs paid by Blue Cross, these rates underestimate the discrepancy. In any case, there is no evidence to support the notion of offsetting trends between commercial and Blue Cross enrollees. However, a subsequent section of this article will provide evidence of important increases in hospital use by small segments of the population *without* private insurance coverage.

While the NHDS data do not separate people with private insurance, there is some fragmentary evidence for the under-65 population from another source. The National Health Interview Survey (NHIS) occasionally provides population-based estimates of hospitalization rates for the insured population. These NHIS data are household-based and thus are not subject to double counting.¶ In 1968, 9.01 per cent of all persons with private hospital insurance aged 64 and under reported a hospitalization; this led to a total of 733 days per 1,000 persons per year.[10] In 1974 the comparable figures were 9.90 per cent and 851 days per 1,000.[11] These data yield an increase similar to that of the NHDS, and the rising use is inconsistent with the BC data, even though BC members represent more than half of the privately insured population.[11]

¶ There are some problems with NHIS data because of nonresponse, forgetfulness, and, most important, the exclusion of data for people dying before the survey is conducted. While these problems will influence the magnitude of the hospitalization rate, there is no reason to expect these biases to affect the trend analyses reported here.

The Duplicate Coverage Problem

The computation of a hospital admission rate obviously depends upon numerators and denominators. The Blue Cross admission rates represent covered inpatient stays divided by Blue Cross enrollment.[12] The AHA and NHDS rates are computed from discharges reported by hospitals and the relevant civilian population (total, or under-65) estimated by the U.S. Bureau of the Census.[4-6] Although one might expect the Blue Cross rates to be more accurate because numerator and denominator are derived from the same organization, consideration of other factors suggests this is not the case. Enrollment figures for Blue Cross, or for any single insuring organization, are not adjusted for the increasing extent of duplicate health insurance coverage in the United States. While carriers may attempt to coordinate benefits when paying claims, they do not count people with multiple policies as fractional enrollees. Therefore, the enrollment figures used in the denominators are overstated. Moreover, duplicate coverage has been rising in the last decade, primarily because an increasing proportion of married women are employed in the paid labor force and because more of these women are working in jobs with employer-sponsored group coverage. A second important factor is that enrollment figures have not adequately tracked the declining size of the American family, again overstating the denominator.

The Magnitude and Growth of Duplicate Coverage

In attempting to clarify the hypothesized problem with the denominator, the situation becomes more complex because there is no simple, accurate way to determine how much duplication truly exists. One method is to use the National Health Interview Survey to ask people whether they have more than one policy for hospital coverage. The proportion responding affirmatively rose from 11.4 per cent in 1968 to 13.0 per cent in 1974, with 0.8 per cent of the latter reporting three or more plans.[10,11] While this indicates that duplicate coverage is real and growing, it probably understates the actual increase because of biases inherent in household surveys. Survey estimates of coverage may be biased downwards because people may not know or may forget that they have coverage.[13–15] It seems even more likely that people will forget *how many* plans provide coverage. This suggests that, as duplicate coverage becomes more common, it may be increasingly underestimated by surveys unless the survey is carefully designed to track down all types of coverage.[16]

Alternatives to survey estimates are provided by the Health Insurance Institute (HII).[8] The HII figures are derived from reports by carriers of the number of people covered under various policies, with adjustments to remove what HII estimates is duplicate coverage. The HII coverage figures are biased upwards because this adjustment is probably incomplete and because of the lag between the time someone leaves a job and loses coverage, and the time the insurer adjusts its records. This results in phantom enrollments—people whose ineligibility is discovered only if a claim is submitted.**

Although the choice between the HII net coverage and survey estimates is important for determining the number of people without coverage, the estimate of *changes* in the level of duplicate coverage is rather insensitive to the choice of data series. Table 1 presents various estimates of the trends in gross and net hospital insurance coverage. The periodic surveys of the National Center for Health Statistics provide unduplicated counts of the number of people with private coverage. There has been some increase in the percentage of the population with coverage—it rose from 75.9 per cent in 1968 to 76.8 per cent in 1976. This increase of less than 1 per cent is small in contrast to year-to-year variations of greater magnitude reflecting either changes in the questionnaire (1972) or substantial increases in unemployment (1975). To minimize the influence of such year-to-year variations, annual rates of increase are computed for each of the data series.

The gross enrollment figures published by the Health Insurance Institute are more problematic because of substantial changes in reporting. The annual HII data include the number of people with four different sources of hospital insurance: 1) individually sold policies written by insurance companies; 2) group policies from insurance companies; 3) Blue Cross–Blue Shield; and 4) independent and other plans, largely employer and union programs and HMOs. HII attempts to remove duplicate coverage *within* each source, but it is unclear how this is done. For instance, the HII estimate for gross enrollment in 1973 was nearly 30 million lower than for 1972, reflecting an adjustment in the esti-

** The HII undertakes periodic samples of claims to determine the proportion of people who say they have duplicate coverage, and thus are eligible for coordination of benefits. Since there are obvious reasons for not reporting duplicate coverage, one can assume that such surveys underestimate duplication, much like tax returns underestimate incomes. The HII adjusts upwards these duplication estimates to account for the underreporting, but these adjustments are not based on hard figures. The fact that the HII periodically reduces retroactively its estimates of the net number of people with coverage supports the notion that their coverage figures have an upward bias.

mate for individual coverage. When the HII makes such a change, estimates of coverage for the total population in earlier years are recomputed, but this retroactive adjustment is not made for the population under age 65. Thus, the series for that population group exhibits discontinuities, particularly when individual policies are included (column 5).

The duplication rate is defined as gross enrollment divided by net enrollment. Thus, a rate of 1.25 means that enrollments reported by carriers are 25 per cent above the number of people with coverage reported in household surveys. This implies that their admission rates will have denominators 25 per cent larger than they should be. The annual rate of increase in the duplication rate is equal to the difference in the annual rates of increase between the numerator and denominator—that is, the HII and NCHS figures, respectively.††

Because there are several sets of figures, there is potential for controversy in the choice of data.‡‡ Fortunately, the results using various reasonable alternatives are all quite close. Ideally, the duplication rate would be measured just for people with regular group Blue Cross coverage. (Note that this would be in the denominator; the numerator would include all the policies held by those people.) Such data do not exist. The closest approximation is the NCHS series of net coverage for the under-65 population, column 2 in Table 1. The rate of increase in such coverage, 0.68 per cent per year, may be compared with the increase in gross enrollments for the under-65 population, columns 5 and 6. Enrollment under "all but nongroup commercial policies" for these people (column 6) has been increasing at 1.70 per cent per year, or 1 per cent more rapidly.§§ (Column 6 – Column 2) As noted, a consistent set of data is not available for all policies held by the under-65 population (column 5), so an annual rate of change cannot be meaningfully computed. The two HII series for all age groups, columns 3 and 4, exhibit remarkably similar rates of increase: 2.08 per cent and 2.21 per cent. Moreover, the annual increase in duplicate coverage for all persons (column 3 or 4 minus column 1) is also about 1 per cent per year, regardless of the HII series used.

Thus, the best estimates of the annual rates of increase in reported enrollments exceed true coverage increases by about 1 per cent per year. The enrollment figures are crucial to the calculation of admission rates; therefore this duplicate coverage problem accounts for half the observed 2 per cent difference between admission-rate trends reported by Blue Cross and the National Hospital Discharge Survey.

The preceding analysis rests on the crucial assumption that the growth rate in dup-

†† The derivation of this is as follows: let G and N be the gross and net enrollment levels at a point in time, so G/N is the duplication rate. The rate of increase in G/N is $\frac{d\,(G/N)}{dt} = \frac{dG}{dt} - \frac{dN}{dt}$. Since G and N can be expressed as $G = a_0e^{b_0t}$ and $N = a_1e^{b_1t}$, $\frac{dG}{dt} = b_0$, $\frac{dN}{dt} = b_1$. Thus, $\frac{d\,(G/N)}{dt} = b_0 - b_1$.

‡‡ In fact, the numbers *within* each column are open to question. The estimates from NCHS of the unduplicated persons with coverage (column 2) are subject to random sampling error and an adjustment for unknown coverage status. (The latter is a proportionate allocation based on people with known coverage.) Similarly, the HII figures are only as good as their underlying sources. Some Blue Cross plans take special efforts to discourage duplicate BC coverage while others purposely double-count people with known duplicate coverage.[17,18] While problems undoubtedly exist, one cannot determine their likely impact.

§§ This category is the sum of the following three types of coverage reported by the HII: 1) group policies written by commercial insurance companies, including Administrative Service Only agreements and Minimum Premium Plans; 2) Blue Cross–Blue Shield and Medical Society Plans; and 3) Other Plans. Excluded are individual and family policies written on a nongroup basis by commercial carriers.

TABLE 1. Gross and Net Enrollment in Private Health Insurance Plans 1968–77 (000 omitted)

	Net Enrollment—Unduplicated Number Persons with Coverage		Gross Enrollment			
			All Ages		Under Age 65	
	All Ages (1)	Under 65 (2)	All Policies* (3)	All But Nongroup Commercial Policies* (4)	All Policies† (5)	All But Nongroup Commercial Policies† (6)
Year						
1968	150,888‡	141,572‡§	176,647	151,369	183,626	147,175
1969	NA	NA	184,289	158,367	191,655	154,034
1970	154,063‡	143,611‡¶	190,349	163,691	198,590	158,995
1971	NA	NA	193,498	165,725	203,156	160,567
1972	155,253‡#	143,309‡#	196,796	167,785	205,720	161,634
1973	NA	NA	200,710	173,166	186,285	163,568
1974	161,104#	148,349#	206,646	177,839	192,093	168,725
1975	162,378¶	147,295**	212,828	182,713	196,668	172,428
1976	164,235¶	149,643¶	209,648	182,652	192,247	171,510
1977	NA	NA	213,930	185,243	195,013	173,384
Time trend regressions††						
a = Intercept	148,821	140,281	176,933	151,183		148,506
b = Annual rate of increase	0.0109	0.0068	0.0208	0.0221		0.0170
R^2	0.9707	0.8834	0.9533	0.9684		0.9340

NA: not applicable.

* Health Insurance Institute, Source book of health insurance data, 1978–79. Washington, D.C.: Health Insurance Institute, 1979.

† Health Insurance Institute, Source book of health insurance data. Washington, D.C.: Health Insurance Institute, yearly, 1968–79. For the years 1968–72 the enrollment figures for persons under age 65 with all types of policies (column 5) exceed the corresponding figures for all persons (column 3) because the latter figures have been retroactively adjusted to reflect new HII estimates of nongroup coverage. This retroactive adjustment is not available for the population under age 65. Each volume of the Source Book includes data for both the preceding year and revised figures for the year before that. With the exception of 1977, these data are the revised figures.

licate coverage for Blue Cross is equivalent to that for all private health insurance. (The absolute level of duplication could be higher or lower; only the rate of change needs to be the same.) Direct tests of this assumption are not available, but two bits of evidence provide some support. For 1968-77 the growth rate in enrollment in Blue Cross plans is 2.41 per cent per year, in contrast to 1.85 per cent for commercial insurers.[8,9,‡‡] If duplicate coverage were growing *more* rapidly among commercial carriers, their gross enrollment would probably also be growing more rapidly than that of Blue Cross. The second piece of evidence is drawn from an analysis of cross-state variations in duplicate coverage.[19] After controlling for several factors that explain duplicate coverage, such as occupational mix, Blue Cross market share within the state had no impact on the ratio of gross to net coverage in the state. If Blue Cross plans were more effective in eliminating duplicate coverage, a negative coefficient would be apparent in the regression. In fact, the coefficient was positive and insignificant.

Explanations for the Increasing Rate of Duplicate Coverage

Several factors influence the rate of duplicate coverage for people with group policies.¶¶ In many married couples both husband and wife may cover each other. Such duplication may be intentional— to improve coverage through double claims or coordination of benefits. But duplication may also be involuntary. More than two

‡‡ The annual growth rate for commercial insurers, 1.85 per cent, refers to group policies only, so as to avoid the problems with individual policies. The growth rate for individual plus group coverage was only 1.71 per cent.

¶¶ People may supplement group policies with individual coverage. However, the group policy will usually provide primary benefits for hospitalization, so such duplication will have a small influence on estimated utilization rates under group policies.

‖ Mueller MS. Private health insurance in 1973: a review of coverage, enrollment and financial experience. Soc Sec Bull 1975;38(2):21–40.

§ U.S. National Center for Health Statistics. Hospital and surgical insurance coverage: United States—1968. Vital and health statistics, Series 10, number 66. Washington, D.C.: U.S. Government Printing Office, 1972. (DHEW publication no. (HSM) 72-1033.

* U.S. National Center for Health Statistics. Monthly vital statistics 21:9 (supplement), December 18, 1972.

U.S. National Center for Health Statistics. Hospital and surgical insurance coverage, United States—1974. Vital and health statistics, Series 10, number 117. Washington, D.C.: U.S. Government Printing Office, 1977. (DHEW publication no. (HRA) 77-1545.) It should be noted that the insurance question for 1972 was worded differently from previous years and was more difficult to process. Therefore, the 1972 estimates are probably biased downwards. See page 4 in this reference.

¶ Carroll MS. Private health insurance plans in 1976: an evaluation. Social Sec Bull 1978;41(9):3–16.

** U.S. National Center for Health Services Research and National Center for Health Statistics. Health: United States—1978. Washington, D.C.: U.S. Government Printing Office, 1978:409. (DHEW publication no. (PHS) 78-1232.)

‡‡ These regressions are based on the equation $y = a \cdot (1 + b)^t$, where y = annual enrollment, b = annual rate of increase, and t is a time variable, with 1968 = 1, 1977 = 10. The coefficients are estimated after a log transformation, $\log y = \log a + t \log (1 + b)$.

thirds of the workers in private industry with hospital benefits have them paid exclusively by employer contribution.[20] Since the employee usually does not have the option of transferring unneeded duplicate health insurance premiums into other fringe benefits, there is no incentive to decline such coverage. In other instances, duplication may provide valuable complementarity, such as belonging to both a Health Maintenance Organization for first-dollar coverage of ambulatory services and to a conventional insurance plan to maintain coverage for second opinions. If coverage is heavily subsidized such duplication may be worth a small additional premium.

Duplicate group coverage also occurs in ways other than the multi-earner family. A substantial number of people (nearly 4 million in 1976) hold multiple jobs.[5] Although most employer-sponsored health insurance is available only to employees working half-time or more, some multiple job holders are likely to be eligible for such coverage. Others, such as construction workers, must accrue a specified number of hours per time period, such as 320–399 hours per 6 months, to maintain coverage.[21] With such relatively low coverage criteria, it may be possible to combine part-time employment with a full-time job elsewhere.

An administrative form of duplicate coverage occurs through job turnover. Many employees retain coverage for a specified period after leaving an employer. In some cases this eligibility may overlap the coverage obtained from a new employer. (However, some employers have waiting periods for new employees.) Even if the grace period has expired, the carrier may not check the employee's eligibility until the next contract period unless a claim is filed. Thus, there are a number of phantom enrollees through reporting lags.

Social and economic changes in the last decade have altered these factors in ways that increase the duplicate coverage rate. The proportion of married women in the labor force has been steadily rising. It was 47.6 per cent in 1970 and 54.8 per cent in 1976, an annual increase of nearly 2.5 per cent.[5] Among full-time employed workers in 1972, 76 per cent of all married men and 57 per cent of all married women had group coverage through their employers. More important, of those with group coverage, 90 per cent of the married men covered their dependents, as did 63 per cent of the married women.[22] Women have also been changing their occupations—moving from those with relatively poor health insurance coverage to those with better coverage. For instance, using the coverage rates in 1972 for each occupation, the shift in women's job mix from 1965 to 1976 would have increased their group coverage rate from 56 to 59 per cent.[22]

Throughout the 1960s and 1970s employers have been increasing their contributions to fringe benefits more rapidly than wages. Part of this increase has taken the form of eliminating the employee contribution for health insurance.[23] Another aspect is the increasing number of employer plans that provide complete coverage for dependents without additional employee contribution. Both trends will increase the duplication rate.

Other administrative factors may also explain part of the increase in duplicate coverage. The unemployment rate among married men had markedly increased, from 2.1 per cent in 1967-68 to 5–6 per cent in 1975-76.[24] This suggests more phantom enrollees because of job changing. The problem of inaccurate reporting may be exacerbated by the rapid rise in Administrative Services Only (ASO) and "self-certification" arrangements whereby firms self-insure and use carriers primarily to handle claims and benefits paperwork.[8] Carriers involved in such arrangement have even less incentive to keep careful track of enrollment changes and in some instances do not keep enrollment files, in-

stead forwarding claims to the employer to determine eligibility.

A final aspect of the duplicate coverage problem arises from the fact that few carriers have an actual head count of their enrollees, irrespective of their enrollees' other coverages. Instead, carriers know how many contracts are in force, and whether they cover one person or a family. (Some family contracts are further classified by two-person and three-or-more-person coverages.) Occasionally, surveys may be undertaken to determine how many people are covered under the average family contract, but such surveys depend upon the carrier and are infrequent. The ratio of persons covered to primary insureds (i.e., contracts) was constant at 2.459 from 1965 to 1970 and fell slightly to 2.423 by 1975.[8] In contrast, average U.S. family size fell from 3.70 in 1965 to 3.39 in 1976.[5] More important, the ratio of persons aged 0–64 to workers aged 16–64 fell from 2.374 in 1965 to 2.017 in 1975, a 15 per cent decline, in contrast to the 1.6 per cent decline in coverage ratio reported by the carriers.[25] This lag in adjusting the family factor leads to an overestimate of the number of persons having insurance.

There have also been marked changes in family structure and enrollment structure that are hidden by the aggregate statistics. The number of households headed by women with children has increased 62.3 per cent from 1970 to 1977.[27] Since children are low utilizers of hospital care, this change in enrollee mix may have an impact on measured use. For some plans, the proportion of single contracts had been rising, which implies that the relatively constant ratio of enrollees/contract reported by the HII is even more biased for family coverage.

Additional Factors Potentially Explaining the Diverging Trends in Hospital Admission Rates

The estimated duplicate coverage effect accounts for only half the differential in hospital admission rates. Part of this shortfall may be attributed to the fact that the difference in utilization trends is based on Blue Cross figures while the duplicate coverage rates refer to all insurers. If commercial carriers had more accurate counts of their enrollees, either through more realistic family factors or fewer phantom enrollees arising from job changes, then the true duplicate coverage rate for Blue Cross enrollees would be higher than for all enrollees, and thus would explain a larger fraction of the utilization difference. Employees with Blue Cross coverage tend to have less restrictive eligibility requirements than those with commercial coverage and this may extend to post-employment coverage as well.[21] On the other hand, if Blue Cross plans had more accurate counts of their enrollees, then even more of the differential is unexplained.*

Blue Cross plans tend to offer more extensive coverage[20] but commercial carriers have been expanding their benefits over the 10-year period and have increased the proportion of actual days paid by the plan.[8] This benefit expansion could have resulted in increased utilization and, more important, such an increase might be expected and thus not a source of concern. But this explanation does not fit the more rapid growth in hospital benefits per enrollee experienced by Blue Cross enrollees. At the same time, improved coordination of benefits among plans may have reduced the number of hospital days paid for twice. This would reduce the BC numerator, although not the actual number of days reported by the AHA or the NHDS.

There is also evidence of increased hospital utilization by people not covered by

* However, as the previously discussed evidence from enrollment statistics suggests, the Blue Cross duplicate coverage rate was probably growing more, rather than less, rapidly than that for commercial carriers.

private health insurance.† Since the utilization differentials appear for the population under 65 years of age, we need not be concerned with Medicare use by the elderly. However, effective in 1973 Medicare coverage was extended to disabled beneficiaries receiving Social Security cash benefits. This could have had a twofold effect. First, the disabled tend to have a high need for medical care, yet in 1972 only 38 per cent had private hospital coverage.[27] The new Medicare coverage may have resulted in an increased use of hospital services, increasing the AHA/HDS figures without altering the BC patterns. Second, Medicare coverage is always treated as primary, so even if the disabled retained their Blue Cross policies, the hospitalization would be attributed to Medicare. While the utilization rate for such people may not have increased substantially, the effective transfer of such high users out of Blue Cross would lower the average BC rate.

The number of hospital days for Medicare disabled beneficiaries rose from 2.5 million in the second half of 1973 to 9.6 million in all of 1977. Discharges rose from 233,020 in 1973 to 937,245 in 1977.[28] The use by Medicare disabled beneficiaries represents about 10 per cent of all patient days attributable to the under-65 population.[29] People covered under the End Stage Renal Disease Program account for about 6 per cent of the discharges of these Medicare disabled beneficiaries. While this is substantial, we cannot assume that none of these patients received treatment prior to Medicare coverage, nor that all switched from Blue Cross coverage. Some must have been covered by commercial insurers and many had no coverage at all. Moreover, the divergence in BC and NHDS utilization is evident prior to 1973 (Fig. 3.)

The second major source of public funding is the Medicaid program. Unfortunately, national statistics on hospital admissions and days are not available. Table 2 presents an estimate of Medicaid admission rates derived from expenditures on inpatient services in general hospitals and the average expense per admission. Although these are only rough estimates, they are of the right order of magnitude. The figures indicate a moderate increase of 2.3 admissions per 1,000 population between 1968 and 1972, with a small decline thereafter. The comparable increase for the same period in the NHDS was 8.3 admissions per 1,000 population.

A third potential explanation for the differential is the increasing use of general hospitals for psychiatric care. Among the population under age 65, the number of discharges with diagnosis of mental disorders per 1,000 population rose from 5.38 in 1971 to 7.54 in 1977.[6] Almost all group insurance plans provide coverage for confinement in general hospitals for mental disorders, so a major portion of this increase is probably covered by Blue Cross and thus cannot account for the differential.[20] The increasing role of federal- and state-funded community mental health centers may have had an influence on the increased use by the noninsured population.[30]

What does all this mean with respect to interpreting the different hospitalization trends reported by Blue Cross and the NHDS? While Blue Cross incentives and controls may have had a substantial differential impact on hospital use, it seems premature to attribute all the divergence in trends to such policies.‡ (Note that to ex-

† More precisely, there was increased use by people with public coverage through Medicare and Medicaid. Such people may or may not have private coverage. Utilization of federal hospitals, such as Veterans Administration facilities, is excluded from both the NHDS and BC figures so changes in such use should not alter the analysis.

‡ Moreover, local cost-control innovations that may be supported by logic and initial evaluations sometimes have little or no overall effect. For instance, recent evaluations suggest that a second surgical opinion program has actually increased surgery rates.[31]

TABLE 2. Estimated Use of General Hospitals by Medicaid Beneficiaries—1968–1977

Fiscal Year	Medicaid Inpatient Services in General Hospitals*† (in Millions)	Average Expense Per Adjusted‡ Admission	Estimated Medicaid Admissions (Thousands)	Estimated Medicaid Admissions Per 1,000 Persons Under Age 65§
1968	$1,361	$471.30	2,888	16.3
1969	1,586	539.25	2,941	16.5
1970	1,887	610.10	3,093	17.1
1971	2,288	675.01	3,390	18.5
1972	2,557	744.88	3,433	18.6
1973	2,660	796.65	3,339	18.0
1974	2,887	878.84	3,285	17.6
1975	3,411	1016.79	3,355	17.9
1976	3,938	1168.15	3,371	17.9
1977†‖	3,678	1316.05	2,794	14.7

* U.S. House of Representatives. Committee on interstate and foreign commerce. Data on the Medicaid programs, eligibility, services, expenditures, fiscal years 1966–77. Washington, D.C.: U.S. Government Printing Office, 1977:32 (years 1968-71).

† U.S. Health Care Financing Administration. HCFA program statistics. Health care financing review 1979:1(2):74. 1972–76 are final figures; 1977 is preliminary.

‡ U.S. NCHS and NCHSR. Health: United States—1978. Washington, D.C.: U.S. Government Printing Office, 1978:416. (DHEW publication no. (PHS) 78-1232.)

§ Civilian noninstitutionalized population under age 65 from U.S. NCHS. Current estimates, yearly.

‖ Expenditures adjusted for the 15-month fiscal year ending September 1977.

plain the differences in trends, the policies must have had a differential impact on BC patients. For instance, for the same medical problem, a physician would have to be more reluctant to admit a patient with Blue Cross coverage than one with, say, Aetna or Prudential coverage.) Instead, about half the difference in trends may be explained by the increasing rate of duplicate coverage, and thus reflect a data artifact. Other parts of the difference may be explained by increasing public coverage through Medicare and Medicaid and publicly funded community mental health centers. While people with public coverage are a relatively small fraction of the population, they represent a disproportionate share of hospital use. Thus, it is not really appropriate to compare the use of people with private coverage to the totals in the HDS and AHA data. Once all these other nonquantifiable factors are considered, it is difficult to determine the true magnitude of the Blue Cross policy changes.

Implications for Policy and Analysis

The increasing importance of duplicate health insurance coverage has implications for both policy and analysis. The policy implications relate to rate setting and to the design of certain approaches to financing health insurance. The analytic problems arise for both academic researchers and in-house analysts monitoring plan performance.

Prospectively determined insurance rates are usually based upon current utilization rates and projected costs. To the extent that utilization per enrollee falls because of duplicate coverage, projected premiums will exceed actual costs. Although the utilization trend for BC members is clearly downward sloping, an ac-

tuary would conservatively use the most recent estimate rather than assume a continuation of the trend. If all premiums are effectively experience-rated, the surplus will be returned. One may question, however, whether true cost-containment measures are more likely to be undertaken in situations of artificial surpluses or in situations of real budget constraints. Surpluses may increase the ability to finance new adminstrative controls, but the tightness of such controls probably reflects budget pressures.

Even if all premiums are experience-rated, the duplicate coverage problem may lead to inequities based upon factors other than utilization differences. The cost of health insurance per enrollee will be lower among employees in firms with a high proportion of working spouses. (In some ways this is analogous to the "marriage tax" incorporated in the federal tax laws.) More important, the second earner in a family gets relatively little from his or her employer contribution for health insurance—that fringe benefit actually accrues to others through lower premiums. This situation would be altered radically by allowing workers to shift the employer contribution to other types of fringe benefits. For instance, if one spouse chose family health insurance coverage, the other might use the comparable employer contribution for a retirement plan. Obviously, such a shift might increase costs to the employer. Under the present system, a firm with a large proportion of married workers might save a substantial amount by tailoring health benefits so that married people would choose to be covered under their spouse's plan, thereby lowering the experience-rated costs to the firm. Any scheme for financing national health insurance through employer contributions must address the situation of multiearner families.

The findings of this article also raise serious problems for analysts. The experienced researcher knows that data are not always what they seem to be at first glance. Yet, finding problems with such crucial figures as hospitalization rates should give one pause. For instance, much of the discussion of the competitive impacts of HMOs has analyzed declines in hospital use by Blue Cross enrollees or lower hospital use by Blue Cross enrollees in areas with high HMO penetration.[32–34] Yet duplicate coverage may be directly related to HMO enrollment for two reasons. Historically, HMOs have developed in areas with extensive employer contributions, making duplicate coverage less expensive to the enrollee. In some cases, HMO and conventional coverage tend to complement one another, making such duplicate coverage more desirable than duplicate conventional coverage.§

Problems also arise for the monitoring of cost-containment activities of insurers. For instance, it becomes much more difficult for a plan to evaluate a preadmission screening program, ambulatory surgical coverage, or outpatient benefits if there are concurrent, unmeasured changes in the local duplicate coverage rate.[3] Since duplicate coverage is highly dependent on local occupational and labor market factors, one

§ To give an extreme example, the health insurance market in Hawaii is dominated by Kaiser and Hawaii Medical Services Association (HMSA). HMSA coverage typically includes some copayments and if someone has multiple HMSA coverages, only the most comprehensive plan applies; coordination of benefits is not used to provide 100 per cent coverage. Hawaii has an extremely high rate of multiearner families, many of whom have dual choice (Kaiser or HMSA) options. Since multiple HMSA coverage provides no additional benefits, one might anticipate a very high rate of duplicate coverage. This may explain the rapidly falling hospitalization rates for both Kaiser and HMSA in spite of an increasing hospitalization rate reported by the AHA. The decision to have HMO plus conventional coverage is actually very complex. It depends on the types of plans involved, the family's usual sources of care and the net premium costs. While a clear understanding of these enrollment factors is not yet available, the potential for duplicate coverage must be considered.

cannot easily apply national trends or use "matched" controls.[19]

Having recognized the problem, is it possible to do something about it? If policy changes are undertaken to encourage the flexible use of fringe benefits for things other than health insurance, the problem may disappear. The identification of a statistical artifact is unlikely to have such a policy impact, but the recognition that duplicate coverage is a large and growing factor might help efforts in that direction. While the problem remains with us, the National Center for Health Statistics might attempt to measure more accurately the extent of multiple coverage. At a minimum, insurers could attempt to maintain more accurate head counts of the number of eligible enrollees. Some carriers already require the use of a unique identifier, such as the Social Security number. Alternatives could be designed to protect individual privacy while allowing computerized matching of people covered by multiple carriers. If carriers know how many people have multiple coverage, they can then adjust the enrollment statistics accordingly. Until such changes are made, however, users of such data should interpret utilization rates with extreme caution.

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Mode of Payment as a Predictor of Health Status, Use of Health Services and Preventive Health Behavior:

A Report From the Los Angeles Health Survey*

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Several issues relevant to Health Maintenance Organizations (HMOs) were examined in this article using data collected from the Los Angeles Health Survey. No support was found for the hypothesis that HMOs disproportionately attract people in poorer health. In fact, HMO members actually reported lower rates of acute or occasional illness and disability than fee-for-service subscribers (FFS). HMO members also reported lower total family incomes and educational status than FFS subscribers and they were less likely to report having a regular doctor. Few differences were found between the two groups in reports of using health services and practicing preventive health behavior. However, HMO members were more likely to report a recent physical examination—but only after taking into account the fact that they were less likely to have a regular doctor. It has been shown that these findings are consistent with recent evidence, and serve to amplify several key findings reported in other studies.

THE CONTINUING ESCALATION of health care expenditures is one of the most serious and controversial issues facing policymakers in the health care field. During the past decade expenditures for health care have increased at the rate of 12 per cent per annum, resulting in a *twofold increase* in health care costs within a 6-year period (1970–1976).[1,2] One of the advantages claimed by HMOs (particularly closed-panel HMOs) is that they will encourage reductions in the costs of medical care, especially costs associated with hospitalization. Although previous research is consistent with this hypothesis, different interpretations have been offered to explain the lower hospitalization rates among HMO members.[3–8] Advocates of HMOs argue that physicians in a prepaid group practice will be less likely to prescribe unnecessary medical treatments, since their fee is not dependent on the number of treatments they prescribe. In addition, since HMO physicians have a financial incentive for maintaining the health of their clients, it is believed they will encourage more preventive health care than FFS physicians.[3,5,6–9] Critics of HMO programs, on the other hand, claim that HMOs have a latent financial incentive to undertreat patients. Others have noted that the lack of a close doctor–patient relationship

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among members of a prepaid group practice may inhibit use of health services.[10] It has also been suggested that prepaid health insurance will attract people in greater need of health care, resulting in an "adverse self-selection bias" in HMOs.[11]

In this article HMO members have been compared with FFS subscribers on a number of measures related to these issues, including self-reports of illness and disability, preventive health behavior and use of physician services. These data were collected by personal interview as part of the Los Angeles Health Survey. Thus, they reflect the health care experiences of HMO and FFS members residing within a large, diverse metropolitan area. As will be shown below, our results corroborate and amplify findings from several recent studies.

Methods and Procedures

The Los Angeles Health Survey (LAHS) is a longitudinal survey of the general population of Los Angeles County that was initiated in 1974 by the late Leo G. Reeder. The overall research design incorporates both a trend and a panel survey. In the trend survey we have been comparing independent samples drawn from the same population at different points in time.[12] The main objective of the trend survey is to monitor aggregate population trends in health status, use of health services and adherence to medical advice. The panel survey permits analyses of these same variables using the individual (rather than the group) as the unit of analysis. In the panel survey respondents were reinterviewed every 6 to 8 weeks for approximately 1 year (1976–1977). The data reported below were obtained from the first panel interview conducted during the fall/winter of 1976. Analysis of the longitudinal date is currently in progress, and will be reported in a subsequent article.

The sampling design used in this survey is a three-stage random probability sample of Los Angeles County that was developed by the Institute for Social Science Research (ISSR) at UCLA.[13] The sampling frame contains approximately 20,000 computer-readable addresses sampled from the county on an area-probability basis. Samples drawn from the frame may be characterized as probabilities proportional to size (PPS) three-stage samples. The three stages use a computer sampling procedure which ensures that each housing unit in Los Angeles County has an equal probability of being selected.

Initially, 2,020 housing units were selected for the 1976 sample. However, 137 of these addresses were declared ineligible (e.g., vacancies, no dwelling unit, etc.), resulting in an eligible sample of 1,883. One adult respondent was then randomly selected from each household using the Kish[14] selection table. Of these 1,883 potential respondents, 1,210 (64.3 per cent) were successfully interviewed. Eighteen per cent of the eligible sample refused to participate, while another 10 per cent were coded as not at home (after a minimum of three callbacks). Nonresponse from all other sources totaled 8 per cent.

The LAHS interviews were conducted in the respondent's home, and were about an hour in length. The interview schedule covered a wide range of topics related to the health and health behavior of the respondents, including recent illnesses and disabilities, use of health services for these illnesses and preventive health behavior. Other questions asked about health insurance coverage. During the interview respondents were handed a card listing more than 20 health insurance companies and health plans; they were then asked whether they had health insurance. Those responding yes were asked to name their companies or health plans. This information was subsequently coded into five general categories of health insurance coverage: 1) no insurance; 2) Medi-Cal coverage; 3) Medicare coverage; 4) private fee-

for-service insurance; and 5) prepaid HMO coverage. In the analyses reported below, comparisons have been made between the HMO and FFS groups. The HMO group originally contained 200 respondents. However, 10 of these people also belonged to a FFS health plan. To keep the two groups as distinct as possible, these 10 respondents were deleted from the analysis. Of the remaining 190 people in the HMO group, 81 per cent belonged to Kaiser, while another 14 per cent belonged to Ross–Loos. The FFS group contained 653 respondents, which excludes those belonging to an HMO, as well as those reporting Medicare or Medi-Cal coverage. The most popular insurance plans in the FFS category were Blue Cross (28 per cent) and Blue Shield (9 per cent).

The HMO–FFS comparisons reported in this article focus on the following dependent variables obtained from the household interviews:

Self-reported tendency to delay
Use of physician services for acute conditions (past 2 months)
Use of physician services for chronic conditions (past 2 months)
Physician recommended preventive health behavior
Recency of last reported physical examination
Cessation or reduction in smoking
Breast self-examination

The scoring procedures for these variables, as well as their mean scores, are presented in Table 1. The actual questionnaire items used to create these variables are presented in Appendix A.

The nine dependent variables shown in Table 1 were examined using multiple regression. The "mode of payment" variable was coded in dummy variable format

TABLE 1. Variable Names, Scoring Procedures and Mean Scores for the Dependent Variables

Variables	Scoring	$\bar{X}$	N*
Self-reported tendency to delay	1 = respondent says he/she is a "long delayer" 0 = respondent says he/she delays a moderate or short amount of time	0.22	843
Use of physician services for acute conditions	Proportion of acute problems reported by respondent that resulted in doctor contact (past 2 months)	0.51	248
Use of physician services for chronic conditions	Proportion of chronic problems reported by respondent that resulted in doctor contact (past 2 months)	0.32	410
Physician recommended preventive health behavior (PHB)	1 = respondent says physician recommends some form of PHB 0 = respondent reports no PHB is being recommended	0.23	843
General physical examinations (GPE)	Number of years since person last reported a GPE. Scores ranged from 1–8+ years.	2.63	843
Cessation of smoking (among those who smoked for at least 1 year)	1 = respondent says he/she quit smoking cigarettes 0 = respondent reports he/she still smokes	0.34	491
Reduction in smoking (among current smokers)	1 = respondent reports reduction in smoking when compared to "recent years" 0 = no reduction reported	0.29	323
Breast self-examination (BSE) (among women)	1 = respondent says she practices BSE 0 = no BSE is reported	0.61	442
Frequency of BSE	1 = respondent reports BSE monthly or more often 0 = no report of monthly BSE	0.39	442

* Group size will vary depending on the item. The base number of 843 represents the combined categories of HMO (N = 190) and FFS (N = 653).

(HMO = 1; FFS = 0), and entered as one of the predictor variables. Other predictor variables include a host of demographic and health status measures that serve as control variables in this analysis (e.g., age, sex, educational and occupational ranking, total family income, recent illness history, etc.). In the interest of space, we have only reported the regression coefficients obtained for the "mode of payment" variable. However, the regression coefficients for the other predictor variables are available for review in Appendix B.

As shown in Table 1, virtually all dependent variables are scored as either dichotomies or proportions (i.e., the scores ranged from 0 to 1). For such variables, the size of the (unstandardized) regression coefficient for "mode of payment" represents the difference in proportions between the HMO and FFS groups—estimated after all other variables in the equation have been controlled statistically. Given the coding of the mode of payment variable, a positive regression coefficient would indicate that the HMO group scored higher on that dependent variable than the FFS group, while a negative coefficient would indicate the opposite. Thus, if the FFS group scored at the mean for "general tendency to delay" ($\bar{X} = 0.22$), and the regression coefficient for mode of payment was −0.02, then the proportion for the HMO group would be 0.02 lower than the FFS group (or 0.20). Of course, if the regression coefficient had been positive rather than negative, then the HMO group would have scored 0.02 higher (or 0.24).

Results

Before comparing the HMO and FFS groups in their (self-reported) use of health services and preventive health behavior, we should first determine their comparability on other related factors, including differences in sociodemographic characteristics and recent morbidity. As shown in Table 2, HMO members had lower education and total family incomes, and they were more likely to be black and Hispanic. HMO members were also more likely to have children (under the age of 18) living in the household. There were no differences between the two groups in age, sex or occupational ranking.

Given this demographic profile, one might expect HMO members to be at slightly greater risk for health care than FFS subscribers. However, HMO members actually reported less acute illness and disability within 2 months of the interview (see Table 3). There were no differences in reports of chronic illness or in self-rated health status.

The lower rates of acute illness reported by HMO members occurred primarily in symptoms of the respiratory tract. As shown in Table 4, 7.9 per cent of the HMO group reported acute respiratory symptoms during the past 2 months, compared with 16.8 per cent of the FFS group. Differences were most evident in reports of cold and flu symptoms. Eight HMO members reported cold symptoms (4.2 per cent), compared with 56 FFS subscribers (8.6 per cent). Similarly, only five HMO members reported the flu (2.6 per cent), compared with 48 FFS subscribers (7.4 per cent). Aside from these two differences, however, the HMO and FFS groups reported very similar rates of acute illness. The distribution of chronic symptoms also appears comparable, with the largest differences occurring for symptoms of the nervous system and mental health.

Use of Physician Services

Most studies have shown a tendency for HMO members to report lower rates of hospitalization and surgeries,[3–8,15–25] but higher rates of ambulatory and physician visits.[3,4,6–8,26–33] Since our measures of use deal exclusively with physician visits, we expected to find higher reporting levels in the HMO group. Contrary to expectation, however, this hypothesis was not sup-

TABLE 2. Cross-tabulation of Demographic Characteristics by Mode of Payment

	HMO (N = 190)	FFS (N = 653)	p ≤
Sex			
Female	52.7	52.4	NS
Male	47.3	47.6	
Age			
18–29	30.7	33.6	
30–39	24.4	21.8	
40–49	20.1	18.9	NS
50–59	14.9	19.4	
60+	10.0	6.3	
Education			
Less High School	28.3	18.8	
High School	24.6	26.3	
Some college	22.9	31.1	.01
College grad and above	24.1	23.8	
Occupation			
Operatives	12.4	13.0	
Crafts	6.0	10.1	
Sales/clerical	27.2	31.7	NS
Managerial	9.4	10.4	
Professional/technical	20.2	20.9	
Income			
< 5,000	7.4	2.9	
5,000–9,000	16.6	16.2	
10,000–13,999	22.1	19.7	
14,000–19,999	24.4	24.9	.05
20,000–24,999	13.2	14.3	
25,000+	16.3	22.0	
Race			
White/Anglo	59.8	75.2	
Hispanic	22.4	17.6	.001
Black	17.8	7.2	
Children in household			
Yes	56.2	47.8	
No	43.8	52.2	.05

NS: not significant.

ported. To illustrate this point, we have presented the regression coefficients for "mode of payment" in nine separate analyses. As shown in Table 5(A), the HMO and FFS groups did not differ statistically in their self-assessed tendency to delay, or in the proportion of acute and chronic health problems that resulted in a doctor contact within 2 months of the interview.

In Table 5(B) we have presented comparable results obtained for several particular symptoms. These measures were obtained for all symptom codes that satisfied two nominal criteria: the number of cases was ≥ 50, and the distribution of the proportion had no more than 75 per cent of the cases in the modal category. Six symptom codes satisfied these criteria: acute respiratory problems, acute flu symptoms, acute musculoskeletal problems, chronic cardiovascular problems—particularly high blood pressure—and chronic musculoskeletal problems. As shown, HMO members reported fewer doctor contacts for acute respiratory and chronic musculoskeletal symptoms, but more frequent doctor contacts for acute

TABLE 3. Mean Scores for HMO Members and FFS Subscribers in Self-Reports of Illness, Disability and Health Status

	$\bar{X}$ Scores	
	HMO (N = 190)	FFS (N = 653)
Average number of acute illnesses reported past 2 months	0.23	0.39*
Average number of chronic illnesses reported	0.68	0.70
Proportion reporting bed days past 2 months	0.11	0.23*
Proportion reporting reduced activity days past 2 months (not counting bed days)	0.12	0.28*
Self-rated health status (proportion reporting excellent or good health)	0.83	0.86

* $p < .001$ (F-test).

musculoskeletal symptoms and symptoms of the cardiovascular system. However, since these differences are based on a smaller number of cases, the regression coefficients would have to be much larger to attain statistical significance (i.e., none of the regression coefficients reported in Table 5 is statistically different from 0). Thus, the picture that begins to emerge from our data is one of similarity between the two groups in their use of physician services.

This conclusion is reinforced even further when we compare the two groups in terms of their "use/disability ratios." The use/disability ratio is an aggregate index for groups that measures the number of doctor contacts per 100 disability days (bed days + reduced activity days).[12,34,35] Thirty-four people from the HMO group reported 263 disability days and 26 "initial" doctor contacts. Accordingly, there were approximately 9.9 doctor contacts per 100 disability days in the HMO group. In the FFS group 246 people reported 2,084 disability days and 183 doctor contacts, for a use/disability ratio of nearly 8.8 doctor contacts per 100 disability days. In other words, the two groups differed by slightly more than one doctor contact per 100 disability days—a very small difference indeed.‡

Preventive Health Behavior

One frequently hears that prepaid health care will create a financial incentive for the practice of preventive medicine. The available evidence, however, is somewhat inconsistent on this point. Some studies have shown that HMO members receive more preventive health care than FFS subscribers; other studies have found little or no differences between modes of payment.[9] Our results are much more consistent with the latter group of studies. As shown in Table 6, there were no differences in reports of receiving preventive health advice from physicians, or in the recency of physical examinations. Similarly, there were no differences in the percentages of exsmokers or in the percentage of current smokers who report smoking less. Among the women, there were no differences in reports of BSE.

Identification With a Regular Doctor

During the interview all respondents were asked whether "there is one person you usually see" for your health care and whether "this person is a medical doctor."

‡ We might note that the FFS group had a somewhat larger proportion of people reporting long-term disability—particularly people reporting 30 or more restricted activity days during the past 2 months. For example, only 3.2 per cent of the HMO group reported 30+ days of restricted activity, compared with 7.9 per cent in the FFS group. However, the proportion reporting 25+ restricted activity days was about the same for the two groups. To adjust for the effects of these "outliers," we truncated the individual reports of disability at 25+ days and recalculated the use/disability ratios. Interestingly, these adjusted rates were virtually identical for the two groups: 11.45 doctor contacts/100 disability days for the HMO group, compared with 11.74 doctor contacts for the FFS group.

TABLE 4. Frequency and Per Cent Reporting Particular Symptom Codes by Mode of Payment

Symptom Categories*	No. and Per Cent Reported as Acute Symptoms				No. and Per Cent Reported as Chronic Symptoms			
	HMO (N = 190)		FFS (N = 653)		HMO (N = 190)		FFS (N = 653)	
	N	%	N	%	N	%	N	%
Nonspecific (0.0–0.49)	0	0.00	6	0.92	4	2.11	10	1.53
Nervous system (50.0–99.0)	0	0.00	4	0.61	12	6.32	21	3.21
Skin, nails and hair (100–199)	1	0.53	15	2.30	2	1.05	21	3.21
Cardiovascular and lymphatic systems (200–299)	0	0.00	5	0.76	19	10.00	62	9.49
Respiratory system (300–399)	16	7.89	110	16.85	23	12.10	77	11.79
Musculoskeletal system (400–499)	17	8.95	60	9.19	28	14.74	93	14.24
Digestive system (500–599)	6	3.16	23	3.52	11	5.79	42	6.43
Urinary system (600–629)	1	0.53	6	0.92	2	1.05	11	1.68
Reproductive system (630–699)	1	0.53	8	1.23	5	2.63	20	3.06
Eyes and ears (700–799)	3	1.58	4	0.61	7	3.68	19	2.91
Mental health (800–899)	0	0.00	0	0.00	5	2.63	4	0.61

* Symptoms were coded using the National Ambulatory Medical Care Survey Symptom Classification. The numbers in parentheses denote the code values.

Fifty percent of the HMO members reported having a regular doctor, compared with 75 per cent for the FFS group ($p < 0.001$). Recently, Scitovsky and associates[10] have shown that the lack of a close doctor–patient relationship may hinder use of physician services. Thus, one might hypothesize that doctor visits for HMO members will increase, relative to FFS, when differences in having a regular doctor are taken into account. Given the coding for the "mode of payment" variable, this hypothesis would predict that the regression coefficient for mode of payment will increase when "regular doctor" is added to the regression. Interestingly, we found one such case among our dependent variables. As reported earlier in Table 6, when we examined the number of years since the person last reported a physical examination, the regression coefficient for "mode of payment" was −0.22 ($F = 1.32$, not significant). When "regular doctor" was added to the equation, the regression coefficient for mode of payment increased to −0.47 ($F = 5.55$, $p < 0.05$). Thus, we found that HMO members reported a more recent physical examination—but only after taking into account the fact that they were less likely to have a regular doctor.

As noted in Appendix A, use of physician services for a physical examination was analyzed in several different ways. The finding reported in the previous paragraph pertains to the measure based on number of years (trimmed at 8+ years to reduce skewness). When we examined the logarithm of the original variable, the same exact pattern emerged. That is, before adding "regular doctor" to the regression, the regression coefficient for mode of payment was only 0.04 ($F = 0.40$, not significant). However, when regular doctor was added as a control variable, the regression coeffi-

TABLE 5. Regression Coefficients (β) and Sample Size (N) for "Mode of Payment" and Self-Reported Use of Health Services*

	β	N
A. Use of health services—general measures		
Self-reported tendency to delay (proportion reporting being long delayers)	−0.02	(840)
Proportion of acute problems (past 2 months) resulting in doctor contact	0.00	(247)
Proportion of chronic problems (past 2 months) resulting in doctor contact	−0.05	(409)
B. Use of health services—particular symptoms†		
Proportion of acute respiratory problems (300–399) resulting in doctor contact (past 2 months)	−0.07	(125)
Proportion of acute flu symptoms (313) resulting in doctor contact (past 2 months)	−0.14	(53)
Proportion of acute musculoskeletal problems (400–499) resulting in doctor contact (past 2 months)	0.05	(71)
Proportion of cardiovascular problems (200–299) resulting in doctor contact (past 2 months)	0.01	(81)
Proportion of high blood pressure problems (205) resulting in doctor contact (past 2 months)	0.14	(51)
Proportion of chronic musculoskeletal problems (400–499) resulting in doctor contact (past 2 months)	−0.13	(121)

* The "mode of payment" variable was coded 1 = HMO, 0 = FFS. The (unstandardized) regression coefficients represent differences in mean scores between the two groups after taking into account all other variables in the equation, including: respondent's age, sex, total family income, marital status (1 = married, 0 = not married), race/ethnicity (1 = White/Anglo, 0 = Other), educational and occupational ranking, reduced activity and bed days (past 2 months), number of acute illnesses (past 2 months), number of chronic illnesses and self-rated health status. The number in parentheses (N) is the size of the group being analyzed. Group size will vary depending on the item.

† All symptoms were coded using the National Ambulatory Medical Care Survey System Classification. The numbers in parentheses denote the code values.

cient for mode of payment increased to 0.12 ($F = 3.70$, $p < 0.06$). Additional measures included rates of people reporting a physical examination within 1, 2, 3, 4, and 5 years of the interview. For rates that extended 4 years and beyond, the pattern described above was replicated. Thus, when predicting whether the person had a physical examination within 4 years of the interview, the regression coefficient for mode of payment was only 0.04 ($F = 1.46$, not significant). However, after adding "regular doctor" to the regression, the regression coefficient for mode of payment increased to 0.07 ($F = 4.91$, $p < 0.05$). Interestingly, this same effect was *not* in evidence when examining 1-, 2-, and 3-year rates of physical examinations. This pattern suggests that prepaid group practice, when compared with comparable FFS practice, may have a lower proportion of people who delay getting a physical examination for a relatively long period of time (e.g., four or more years). However, there appears to be little difference between the two groups in the proportion of people receiving a more recent physical examination (e.g., within the past 2 or 3 years).

Summary and Discussion of Results

Perhaps our findings can best be organized and discussed in relation to several key policy issues relevant to HMO legislation. For example, one such issue is whether HMOs disproportionately enroll people at "greater risk." In a recent study of this hypothesis, Berki and associates[11,36] distinguished between two forms of

TABLE 6. Regression Coefficients (β) and Sample Size (N) for "Mode of Payment" and Self-Reported Preventive Health Behavior*

Preventive Health Behavior	β	N
Proportion reporting preventive health advice from doctor.	0.01	(839)
No. of years since last reported physical examination. (Scores trimmed at 8+ years).	−0.22	(839)
Proportion stopped smoking (among those who smoked for at least a year).	−0.03	(490)
Proportion smoking less (among current smokers).	−0.03	(322)
Proportion practicing breast self-examination (BSE) (among women)	−0.04	(440)
Proportion practicing monthly BSE (among women).	−0.01	(440)

* The "mode of payment" variable was coded 1 = HMO, 0 = FFS. The (unstandardized) regression coefficients represent differences in mean scores between the two groups after taking into account all other variables in the equation, including respondent's age, sex, total family income, marital status (1 = married, 0 = not married), race/ethnicity (1 = White/Anglo, 0 = Other), educational and occupational ranking, reduced activity and bed days (past 2 months), number of acute illnesses (past 2 months), number of chronic illnesses and self-rated health status. The number in parentheses (N) is the size of the group being analyzed. Group size varies depending on the item.

risk—one related to health, the other to costs (for an earlier treatment of this hypothesis see Bashur and Metzner[37]). According to the financial-risk hypothesis, HMOs will enroll more people vulnerable to financial loss than FFS, resulting in an "adverse economic risk self-selection factor" in HMOs. Data from previous studies are generally consistent with the financial vulnerability hypothesis.[11,26,38,39] Our data also support this proposition, since HMO members had lower total family incomes than FFS subscribers. Additionally, we found that HMO members were more likely to have children living in the household. This finding supports previous research showing that "HMOs attract families expected to exhibit greater demand for care that is associated with a younger and expanding life stage"[11] (p. 607). However, HMO enrollees did not report different rates of chronic illness or poorer overall health. In fact, the only statistical difference we found had HMO members reporting less acute illness and disability than FFS subscribers—particularly cold and flu symptoms. Consequently, we have no evidence to support the health–risk hypothesis that HMOs will enroll more people in poor health. Although a few early studies reported data in support of the health-risk hypothesis,[6,37,40,41] much of the recent evidence is consistent with our findings of no adverse health risk among HMO members.[11,15,26,36,38,39,42,43]

Another key issue is whether the HMO and FFS groups use health sevices in a comparable fashion. Most studies have shown a tendency for HMO members to report higher rates of ambulatory and physician visits than FFS.[3,4,6,18,26–32] However, we found no differences in reports of using physician services for either acute or chronic illness. As noted previously, our indicators of use measure the proportion of health problems that resulted in a doctor contact (within the previous 2 months). Consequently, these measures primarily reflect *entry* into the health care system via a physician visit. They do not reflect use of services for follow-up visits. Thus, if HMO members are more likely to have follow-up visits to the doctor, then this difference would not be detected in this analysis.

One clear difference did emerge between the two groups concerning their identification with a regular doctor. HMO enrollees were much less likely to report a regular doctor than FFS subscribers. Sev-

eral studies have shown a positive association between use of services and having a regular physician.[10,44] For example, Scitovsky and associates[10] have recently compared use of services by Kaiser members with that of a group practice clinic plan that was predominantly fee-for-service.§ They found that the Kaiser sample had lower rates of use than the clinic sample, which they attributed to the smaller percentage of Kaiser members reporting a regular doctor.

The data reported by Scitovsky et al. were obtained from a sample of faculty and staff at Stanford University. Consequently, the generality of their results might be questioned. It is of interest to note, therefore, that we uncovered one finding that is consistent with their results. That is, we found that reports of using physician services for physical examinations increased for HMO members, relative to FFS, when initial differences in having a regular doctor were taken into account. Thus, one might speculate, as Scitovsky and associates[10] have done, that "use of physician services may have been held down by the relative lack of a close patient–physician relationship" (p. 15) among Kaiser members. On the other hand, we should not forget that the overall size of this effect was quite modest in our data. Future research would be well-spent examining this issue in more detail.

Supporters of HMO legislation have maintained that prepaid health insurance will encourage the practice of preventive health care. Although previous research is generally consistent with this proposition,[18,42,45–47] several studies have reported mixed or contradictory results.[19,24,48] Luft[9] has recently reviewed this literature and concluded that "the greater number of preventive services among HMO enrollees seem merely to result from the lower price patients face to ambulatory visits [in HMOs]" (p. 141). In other words, Luft has argued that HMO members receive more preventive health care not because HMO physicians are more committed to the practice of preventive medicine, but because preventive health services are generally less expensive to the consumer in an HMO. Luft's conclusion is particularly noteworthy, since it discounts the popular notion that prepaid health care creates a financial incentive for physicians to practice more preventive medicine. Interestingly, our data tend to corroborate Luft's reading of the literature. For example, if HMO physicians were operating strictly from the perspective of their financial interests (i.e., if they were primarily interested in reducing future treatment costs for the HMO) then we would expect them to offer more preventive health advice than FFS physicians. However, the HMO and FFS groups did not differ in their reports of receiving preventive health advice from their doctors (e.g., stop smoking, exercise more often). Similarly, no differences were found in two health behaviors that might be expected to reduce future treatment costs for the HMO: cancer screening behavior (BSE) and cessation or reduction of smoking. On the other hand, differences were found in a behavior consistent with Luft's hypothesis. Virtually all of the HMO members had prepaid insurance covering "health maintenance examinations," while a much smaller proportion of the FFS group would be expected to have comparable coverage. As Luft would predict, HMO members reported more recent physical examinations than FFS subscribers. However, as noted previously, we observed this difference only after taking into account the fact that HMO members were

§ More specifically, the clinic plan examined by Scitovsky and associates differed from the Kaiser plan in that 1) care was provided by a predominantly fee-for-service group practice whose prepaid patients accounted for only about 15 per cent of its annual gross revenue; 2) the group was not at risk for the hospital costs of their prepaid patients; and 3) the physicians often did not know whether the patient was prepaid or fee-for-service.

less likely to have a regular doctor. This latter finding suggests that HMO members might be more likely to take advantage of their preventive health coverage when they have a close doctor–patient relationship.

As Andersen and Aday[44] note, the importance of having a regular source of care underscores the potential value of HMO legislation. Yet, the advantages associated with closed-panel HMOs (e.g., providing a regular place for health care) may be confounded by certain obstacles that limit or hinder use of services. One such obstacle is the tendency for people without a regular doctor to be enrolled in closed-panel HMOs. What remains unclear about this finding is the extent to which closed-panel HMOs discourage identification with a regular physician, as compared to people simply enrolling in closed-panel HMOs because they have no regular doctor. Several studies have shown that enrollment in closed-panel group practice is more likely to occur among people having no regular doctor.[8,11,26,38,41,43] However, FFS physicians may also have a financial incentive to establish a long-term relationship with their clients that is missing in a closed-panel HMO.[10] Thus, future research might examine whether the absence of a regular doctor is basically an "input" characteristic of the enrollment population of HMOs, or is primarily a "process" variable associated with the formal organization of closed-panel group practice. Such information should be of considerable benefit to health administrators interested in improving the operation and management of existing HMOs.

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Appendix A

All variables examined in this analysis were obtained from the 1976 household interview conducted as part of the Los Angeles Health Survey. The individual questionnaire items used as dependent variables are presented below.

Physical Examinations: "When was the last time you had a general physical examination or check-up (not because of any health problem or injury)?"

Answers to this question were originally coded in either days, weeks, months or years, depending on what was most convenient for the respondent. These responses were then converted into an index measuring the number of years since the respondent reported having a general physical examination. The original codes for this index ranged from 1 to more than 30 years, and were highly skewed. To reduce the skew in this variable, scores were "trimmed" at 8+ years (skew = 1.37). Additionally, the logarithm of this variable was examined (skew = 0.81), as were several dichotomous variables that represent different rates of physical examinations. Five such dichotomies were examined, representing the proportion of people reporting a physical examination within 1, 2, 3, 4 and 5 years of the interview. For the most part, the same results were obtained regardless of which measure was used.

Acute Illness: "Think back to the past 2 months; were you sick or did you have any accidents or injuries at any time during the past two months? Please do not include any recurrent or continuing health problem that you may have all of the time. We are interested only in acute or occasional illness or injuries."

Respondents answering "yes" to this question were asked to describe their acute illnesses (symptomatology and/or diagnosis), and whether they contacted a doctor about each problem. Up to three separate illnesses/injuries were recorded by the interviewers.

Chronic Illness: "Do you have a chronic—that is, a recurring or continuing—health problem, a health problem that you have all the time or makes you ill from time to time?"

The same information was obtained from this question as the one concerning acute illness. That is, respondents were asked to describe their chronic problems (up to three mentions), and the length of time since they last contacted a doctor about each problem.

Proportion of Acute/Chronic Problems Resulting in Doctor Contact

The acute and chronic questions listed above asked whether the respondent had "seen or made contact" with a doctor about each problem that was mentioned, and how long ago this contact occurred. This information was used to create two indexes measuring use of physician services: the proportion of acute problems that resulted in a doctor contact within 2 months of the interview and the proportion of chronic problems that resulted in a doctor contact within the same 2-month period.*

* Use of physician services for chronic illness was also examined using a 6-month and 1-year recall period. All three versions produced comparable results. We have focused on the version using a 2-month recall period because it was also used in several other questions that play a key role in this analysis (e.g., number of bed days and reduced activity days).

Another question asked about the respondent's general tendency to delay using health services:

> "When you think there is something wrong with your health, would you say you:
> Wait as long as possible before seeing a doctor;
> Put off going to the doctor for a short time; or
> Try to see a doctor immediately?"

Two questions asked about restricted activity during the past 2 months:

> *Bed Days:* "During the past 2 months, did you stay in bed all or part of any day because of any illness or injury?"
> *Reduced Activity Days:* "(Besides the days in bed) were there any other days during the past 2 months that you had to cut down on things you usually do because of poor health?"

Respondents were also asked to rate their health on a 4-point scale:

> How would you rate your health overall; would you say your health is:
> excellent;
> good;
> fair; or
> poor?"

Another question asked whether their doctors recommended that they change their behavior or life-style:

> "We have talked about some illnesses and health problems. Now I'd like to find out about any health problems, treatments or medicines that you may be taking or doing. Has a doctor prescribed or recommended that you take medicines, medical treatments or change your way of living in such areas as smoking, special diet, exercise, drinking or some other area?"

Respondents answering "yes" to this question were asked to describe these recommendations. Up to three mentions were recorded by the interviewers. The responses were then coded into 15 categories grouped according to whether they referred to behavioral/life-style changes or medications/treatments and other therapeutic devices (e.g., life-support devices, orthopedic braces, etc.). The specific categories depicting behavioral/life-style changes included: eating or avoiding specified foods or following a special diet, reducing activity, increasing activity or exercise, reducing consumption of cigarettes and alcoholic beverages, changes in work habits and getting more rest/relaxation. The common denominator among these recommendations is that they all exemplify some form of "preventive health advice."

Finally, additional questions asked about smoking behavior and self-examination of the breast. All respondents who reported smoking cigarettes for at least one year were asked whether they still smoke cigarettes, and (if yes), whether they are smoking more, less or about the same amount "as in recent years." The women were also asked whether they examine their brests for lumps, and (if yes) how frequently they perform breast self-examination.

Appendix B

Appendix B contains the scoring procedures and standardized regression coefficients for all predictor variables reported in Tables 5 and 6.

TABLE B1. Scoring Procedures for Selected Predictor Variables

Predictor Variables	Scoring Procedures
"Mode of payment"	1 = HMO; 0 = FFS
Sex	1 = female; 0 = male
Marital status	1 = married; 0 = not married
Socioeconomic index (SEI)	Duncan's Socioeconomic Index, where low score = low SEI, and high score = high SEI. Scores ranged from 001 to 996.
Race/ethnicity (self-reported)	1 = white/Anglo; 0 = other
Income	15-point scale, ranging from <$5,000 to $40,000+
Bed days	1 = 1+ bed days; 0 = no bed days
Reduced activity days	1 = 1+ reduced activity days; 0 = no reduced activity days
Self-rated health status	1 = excellent, 2 = good, 3 = fair, 4 = poor
Number of chronic health problems	1–3+
Number of acute health problems	1–3+

TABLE B2. Standardized Regression Coefficients and Sample Size (N) for Variables used in Predicting Self-Reported Use of Physician Services (General Measures)*

	Dependent Variables—Use of Physician Services		
Predictor Variables	Self-Reported Tendency to Delay (N = 840)	Proportion of Acute Problems Resulting in Doctor Contact (N = 247)	Proportion of Chronic Problems Resulting in Doctor Contact (N = 409)
"Mode of payment"	−0.02	0.00	−0.05
Age	−0.02	0.07	0.08
Sex	−0.05	0.02	0.13†
Marital status	0.04	−0.05	0.00
Socioeconomic index	0.12‡	−0.09	0.00
Race/ethnicity	0.06	−0.06	−0.06
Income	−0.04	−0.04	0.07
Bed days	−0.07	−0.01	0.04
Reduced activity days	0.05	0.03	0.11‡
Self-rated health	0.10‡	0.07	0.20§
Chronic problems	0.01	−0.05	−0.13†
Acute problems	0.04	0.09	0.08
R	0.19	0.22	0.30

* The standardized regression coefficients reported in Table B2 were obtained from the same equations which produced the unstandardized regression coefficients reported in Table 5.
† $p < 0.01$.
‡ $p < 0.05$.
§ $p < 0.001$.

TABLE B3. Standardized Regression Coefficients and Sample Size (N) for Variables Used in Predicting Self-Reported Use of Physician Services (Specific Measures)*

Predictor Variables	Dependent Variables—Use of Physician Services: Acute Respiratory (300–399) (N = 125)	Acute Flu (313) (N = 53)	Acute Musculoskeletal (400–499) (N = 71)	Cardiovascular (200–299) (N = 81)	High Blood Pressure (205) (N = 51)	Chronic Musculoskeletal (400–499) (N = 121)
"Mode of payment"	−0.05	−0.09	0.04	0.02	0.12	−0.13
Age	0.00	−0.18	−0.03	0.09	−0.04	−0.06
Sex	−0.02	−0.11	−0.05	0.11	0.10	0.02
Marital status	0.09	0.12	0.00	−0.05	−0.15	0.03
Socioeconomic status	0.12	0.18	0.06	−0.18	−0.06	−0.08
Race/ethnicity	0.09	0.13	0.04	−0.16	−0.31	0.00
Income	−0.07	−0.03	−0.03	0.13	0.23	0.02
Bed days	−0.06	−0.04	−0.07	−0.21	−0.05	0.08
Reduced activity days	0.04	0.15	−0.06	0.16	0.14	0.14
Self-rated health status	−0.16	−0.07	0.14	0.09	0.20	0.18
R	0.23	0.38	0.16	0.38	0.45	0.30

* The standardized regression coefficients reported in Table B3 were obtained from the same equations that produced the unstandardized regression coefficients reported in Table 5.

TABLE B4. Standardized Regression Coefficients and Sample Size (N) for Variables Used in Predicting Self-Reported Preventive Health Behavior*

Predictor Variables	Dependent Variables: Preventive Health Advice From Doctor (N = 839)	No. of Years Since Physical Exam (N = 839)	Stopped Smoking (N = 490)	Smoking Less (N = 322)	Practice BSE (N = 440)	Practice Monthly BSE (N = 440)
"Mode of payment"	0.01	−0.04	−0.03	−0.03	−0.04	0.00
Age	0.10†	−0.05	0.01	−0.01	−0.03	0.01
Sex	−0.01	−0.15‡	−0.11§	−0.20‡		
Marital status	−0.01	−0.05	−0.06	0.05	0.09	0.04
Socioeconomic status	0.00	−0.04	−0.03	0.10	0.06	0.02
Race/ethnicity	0.06	0.00	0.00	−0.14§	0.01	−0.01
Income	0.01	−0.03	−0.02	−0.05	−0.13§	−0.03
Bed days	0.00	0.04	0.01	0.08	−0.05	−0.07
Reduced activity days	0.05	−0.10†	−0.04	0.07	0.02	0.03
Self-rated health status	0.19‡	0.03	0.11§	0.00	0.00	−0.02
Chronic problems	0.09§	−0.07	0.04	0.02	−0.04	−0.02
Acute problems	0.02	−0.02	−0.01	0.15§	−0.02	−0.05
R	0.30	0.24	0.20	0.31	0.14	0.11

BSE: Breast self-examination.

* The standardized regression coefficients reported in Table B4 were obtained from the same equations that produced the unstandardized regression coefficients reported in Table 6.

† $p < 0.01$.

‡ $p < 0.001$.

§ $p < 0.05$.

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Use of Ambulatory Medical Care by the Poor: Another Look at Equity

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Access to health services by the poor and other disadvantaged groups has improved considerably over the past 15 years. These circumstances have led some to question whether the poor now have equal access to health care. This article presents recent evidence from the 1976–78 National Health Interview Surveys (NHIS) comparing utilization among age, race, and income groups. Without adjustment for health status, the poor have more physician visits than those with higher incomes. After adjusting for age and health status, however, these differences are reversed. Depending on which measure is used, the poor have between 7 per cent and 44 per cent fewer visits than those with income above twice the poverty level. Furthermore, the age- and health-adjusted data show blacks have significantly fewer visits than their white counterparts. In addition, there are large differences among race and income groups in the characteristics of the ambulatory care obtained. Blacks and the poor are much more likely to use hospital clinics and less likely to use private physician offices or telephone consultations. The poor also receive less preventive care. It would appear from the present evidence that still further progress is required to achieve the goal of equity in the distribution of medical care services.

ACCESS to health services by the poor and other disadvantaged groups has improved considerably over the past 15 years. These circumstances have led some to question whether the poor now have equal access to health care. In a recent book, Aday, Andersen and Fleming[1] conclude that while some problems of access persist, utilization of medical services is fairly equitable. Considerable controversy has surrounded this statement. Although most observers agree that gaps in access have narrowed during the past 15 years, substantial disagreement exists on the size, nature and importance of remaining problems.[1–5]

The purpose of this article is to present new results on access to medical care by the poor based on the National Health Interview Surveys for 1976-78. We also place these results in the context of previously published work, especially that of

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The views expressed in this article are those of the authors, and no official endorsement by the Department of Health and Human Services is intended or should be inferred.

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Aday et al.[1] Because the focus is on primary and preventive care, we do not consider use of hospital services.

Data And Methods

Data Sources

The basic data source used is the National Health Interview Survey (NHIS), which is conducted by the National Center for Health Statistics.[6] NHIS is a probability sample of the civilian noninstitutional population with 110,000 respondents in 40,000 households each year. The NHIS data presented in this article were collected during the period 1976 to 1978. Data for all three years were combined in order to minimize sampling error.

Since we will compare our results with those presented by Aday et al.,[1] it is important to note some methodological differences between NHIS and their 1976 survey (carried out by the Center for Health Administration Studies [CHAS] and the National Opinion Research Center [NORC] at the University of Chicago. The CHAS–NORC survey was primarily designed to measure access to medical care. It included 7,787 individuals in 5,432 households but oversampled certain population subgroups believed to have access problems.

The two most salient differences between the surveys are the levels of family income used in analysis and the measurement of physician visits. The NHIS data presented here categorize family income to approximate the official poverty income level adjusted for family size. For example, poverty was defined at $7,000 for a family of four and $4,000 for a family of two. Data published from CHAS–NORC generally use a low-income-category ceiling of $8,000 with no adjustment for family size. The NHIS data result in an estimated 16 per cent of the population below poverty, while the CHAS–NORC low-income group comprises one third of the population. At the other end of the scale, the highest income category used in the NHIS tables (more than 200 per cent of the poverty level) includes 52 per cent of the population, while the CHAS–NORC tables include one third of the population in the highest category.

In defining physician visits, the NHIS data are based on a 2-week recall period while the CHAS–NORC data use a full year recall. Although even the 2-week recall underestimates physician visits, the greater recall bias inherent in the CHAS–NORC method results in a greater underestimate of total physician visits and tends to diminish differentials among population subgroups. The NHIS data include telephone consultations with a physician (excluding calls for making an appointment) while the CHAS–NORC data do not. The inclusion of telephone consultations is particularly important for assessing race and income differentials, since white and high-income persons make more use of the telephone for physician contacts.

The Role of Health Status

Access to health care for any given group can only be assessed relative to the need for such services. Comparisons of utilization among population groups need to consider differentials in health status among those groups. Using a variety of different measures, numerous studies have shown that the poor have worse health status than the nonpoor.[2,3,7]

Health status indicators that measure differences in need for ambulatory medical care are not readily available. Physicians spend much of their time helping people to function more effectively and relieving pain and discomfort. Thus, measures of functional ability and positive well-being are the relevant measures of health status given the objectives of medicine today.

Some promising work in this area has been carried out recently. For example,

Brook et al.[8] have developed measures of physical, mental and social health status that were designed to be sensitive to the utilization of medical care. These measures have been found to be reliable and valid and are based on data from self-administered questionnaires. Unfortunately enough data on these measures are not yet available to compare the poor and nonpoor.

Manning, Newhouse and Ware[9] suggest that self-assessment of health is a good predictor of the probability of using health services and of the quality of services demanded, as well as correlating well with clinical indications of need for care and other measures of health status. Since these are the types of measures available from the NHIS, we shall use them in this study.

A further problem in examining income differentials in health status is the income-depressing effects of illness. That is, illness may lead to loss of work, which in turn reduces income. This phenomenon is especially evident in the age groups 17–64. However, it is relatively absent in the under-17 age group.

Adjusting Utilization for Health Status

Because the poor are less healthy than the nonpoor, it is important to adjust for health status in examining utilization data. Two methods of adjustment will be used in this article. The first method examines the use of physician services after stratifying the total population into subgroups that are similar in their overall health status. We will use the respondent's self-assessment of health to stratify the population into two groups. The limitation of this approach is that even within the strata there may be variation in need. Nonetheless, this method provides a useful measure that can be easily interpreted.

The second method of adjusting for health status uses a ratio, which indicates how a group's utilization compares against some measure of need for services by this group. The most commonly used indicator of this type is the "use disability ratio" (UDR).[10,11] The UDR measures the number of physician visits per 100 disability days for those who have had at least one but no more than 100 disability days.[1] A similar measure will be presented from the 1978 NHIS data. This measure is the ratio of physician visits per 100 bed-disability days. Bed-disability days are used rather than all disability days since we believe the former better represent conditions requiring medical attention and are subject to fewer differences in interpretation across income groups. The "use-bed-disability ratio" (UBDR) includes hospital days, which are excluded from the use-disability ratio. Furthermore, it compares visits to bed-disability days for an entire population subgroup, not just those with at least one bed-disability day.

While useful in summarizing overall patterns, use-disability measures suffer from serious limitations.[3,10] The number of disability days represents only a partial measure of need for physician services. In particular, it is relatively insensitive to differences in disease severity of underlying health status contributing to disability. Furthermore, the indicator assumes that disability days and visits are related in a linear fashion. This assumption omits any consideration of visits that are for preventive purposes, or disabilities that do not require medical care.

Statistical Analysis

Three variables were generally used to cross-classify the data: age (under 17, 17–44, 45–64, 65 and older), race (white and black; other races were omitted due to small sample size) and family income (below poverty level, between poverty and 150 per cent of poverty, between 150 per cent and 200 per cent of poverty, above 200 per cent of poverty; respondents who did

not report income [9 per cent of total] were omitted).

When the data are cross-classified in this detail, many of the individual cell estimates are subject to considerable sampling error. As a result, the values observed in certain cells may not be good estimates of the corresponding true values. A powerful statistical technique for "smoothing" the cell estimates is the generalized linear model fitted by a weighted least-squares algorithm.[12–14] These smoothed estimates "immediately identify for the substantive analyst statistically significant differences while suppressing nonsignificant or purely sampling differences."[14] The results reported in the next section are based on the smoothed estimates from these models. Details of the model fitting procedure and observed (i.e., unsmoothed) data are given in the appendix.

In order to express differences in relative terms (i.e., per cent differences), logarithms of the response measures in each cell (e.g., mean number of doctor visits) were fit to a full linear model. Interaction terms were omitted from the model in a stepwise procedure until only effects that were statistically significant (at the 10 per cent level) remained.

Results

Self-Assessment of Health

NHIS respondents were asked (for themselves and other household members) "compared to other persons _____'s age, would you say that his health is excellent, good, fair or poor?" The estimated proportion reporting fair or poor health by age, race and income is shown in Table 1. The general patterns by age are similar: fewer whites than blacks report fair or poor health. There is also a steady decrease in the proportion reporting fair or poor health as income increases. The effect of income on reported health status is greater for whites than blacks.

Among whites the ratios of the proportion reporting fair or poor health in the lowest to that in the highest income group vary from 2.1 for those 65 and over to 3.7 for the 45–64 age group. For blacks the ratios range from 1.5 for those 65 and over to 2.6 for adults 45–64 years of age. For both races, the largest ratios occur in the two age groups between 17 and 64. This is consistent with the income-depressing effects of illness noted earlier. The two youngest age groups show the largest differences by race.

Other measures of health, such as proportion with limitation of activity due to chronic illness, restricted activity and bed-disability days, exhibit similar patterns. Hospitalization rates and length of hospital stays are also greater for blacks and low-income persons, providing further evidence of more serious illness in these groups.

Physician Visits

The NHIS indicates that without adjustment for age or health status, physician visits for the poor slightly exceed those of higher income people. Those below poverty in 1976-78 made 5.1 physician visits per year while those with incomes over twice the poverty level made 4.9 physician visits.

These overall figures mask significant variations by age and race (Table 2). The aggregate differences in use reflect below-average use by poor children balanced against the higher utilization for poor nonelderly adults. For children under 17, visits increase with income and are 23 per cent lower for blacks than whites. Within each race, those in the two lowest income groups have 26 per cent fewer visits than those in the highest category.

For the 17–44 age group, blacks have 17 per cent more visits than whites. Furthermore, the income differential reverses, with the poor having 37 per cent more visits than the two highest income groups.

TABLE 1. Estimated* Per Cent Reporting Fair or Poor Health by Age, Race and Income, 1976–78

Race and Income	Age Adjusted†	Under 17	17–44	45–64	65 and Older
White					
Below poverty	23.5	7.2	16.9	49.5	42.2
100–150% poverty	16.5	4.6	10.7	35.9	33.3
150–200% poverty	12.6	3.2	8.2	26.6	27.4
Above 200% poverty	7.5	2.5	4.8	13.4	19.9
Black					
Below poverty	28.4	9.6	22.9	55.0	48.7
100–150% poverty	22.6	6.9	16.5	45.5	43.9
150–200% poverty	18.4	5.2	13.5	36.1	38.7
Above 200% poverty	12.8	4.7	9.2	21.1	32.8

* Estimated from generalized log-linear model.
† Age-adjusted by the direct method.

In the two oldest age groups there is no evidence of a race differential. For persons 45–64 years of age, visits decrease steadily with income, with the poor having 42 per cent more visits than the highest income group. For the population 65 and over there are no income differences in utilization.

Thus, poor adults have at least as many physician visits as the nonpoor. Children in families with incomes below 150 per cent of the poverty level, however, have fewer visits than those in higher-income families. These aggregate visit data do not take into account the lower health status among the poor noted in the previous section.

The volume of visits reflects both the proportion of the population that visits a physician and the number of visits made by persons seeing physicians. The first component is a measure of access to the health care system which focuses on patient-initiated visits, while volume includes a substantial proportion of physician-initiated visits for follow-up care. Although

TABLE 2. Estimated* Number of Physician Visits per Person per Year by Age, Race and Income, 1976–78

Race and Income	Age Adjusted†	Under 17	17–44	45–64	65 and Older
White					
Below poverty	5.39	3.66	5.24	7.40	6.69
100–150% poverty	4.95	3.66	4.75	6.25	6.69
150–200% poverty	4.82	4.06	4.51	5.54	6.69
Above 200% poverty	5.00	4.92	4.51	5.22	6.69
Black					
Below poverty	5.37	2.81	5.78	7.40	6.69
100–150% poverty	4.91	2.81	5.24	6.25	6.69
150–200% poverty	4.74	3.12	4.97	5.54	6.69
Above 200% poverty	4.87	3.78	4.97	5.22	6.69

* Estimated from generalized log-linear model.
† Age-adjusted by the direct method.

TABLE 3. Estimated* Per Cent With No Physician Visits in the 2 Years Preceding Interview by Age, Race and Income, 1976–78

Race and Income	Age Adjusted†	Under 17	17–44	45–64	65 and Older
White					
Below poverty	15.6	14.6	14.7	19.2	14.9
100–150% poverty	15.2	13.5	14.7	19.2	14.2
150–200% poverty	13.6	10.5	13.7	17.6	13.4
Above 200% poverty	11.1	7.0	12.2	14.4	11.7
Black					
Below poverty	13.6	14.6	11.9	14.8	14.9
100–150% poverty	13.2	13.5	11.9	14.8	14.2
150–200% poverty	11.7	10.5	11.2	13.5	13.4
Above 200% poverty	9.5	7.0	9.9	11.1	11.7

* Estimated from generalized log-linear model.
† Age-adjusted by the direct method.

both are important components of care, the proportion of the population with no visit may be more affected by barriers to access.

In order to identify relatively serious access problems, we examine the proportion who have not seen a physician within the 2 years preceding interview (Table 3). This proportion is greater among those below the poverty level compared with those above twice the poverty level. The income trend shows in all age groups, especially among whites, but the differential is greatest among children. They are twice as likely as their high-income counterparts to have had no visit in the past 2 years. There are no race differences among children or the elderly, but the proportion with no visit is greater among whites than blacks in the 17–64 age groups.

Physician Visits Stratified by Health Status

Table 4 shows the estimated number of physician visits for respondents reported in fair or poor health. For children under 17 there is a steady increase in visits for both blacks and whites—poor children have 47 per cent fewer visits than children in high-income families. Furthermore, at each in-

TABLE 4. Estimated* Number of Physician Visits per Person per Year for Those Reporting Fair or Poor Health by Age, Race and Income, 1976–78

Race and Income	Age Adjusted†	Under 17	17–44	45–64	65 and Older
White					
Below poverty	9.98	9.14	10.36	10.95	8.90
100–150% poverty	10.68	11.36	10.36	10.95	9.61
150–200% poverty	11.34	13.30	10.36	10.95	10.66
Above 200% poverty	13.51	17.16	12.69	10.95	11.90
Black					
Below poverty	8.97	5.68	11.39	9.39	7.58
100–150% poverty	9.42	7.06	11.39	9.39	8.18
150–200% poverty	9.85	8.26	11.39	9.39	9.08
Above 200% poverty	10.65	10.67	11.39	9.39	10.14

* Estimated from generalized log-linear model.
† Age-adjusted by the direct method.

come level, blacks have 38 per cent fewer visits than whites.

Among blacks age 17–44 who report fair or poor health, there is no statistically significant income differential. For whites in this age group the three lowest income groups have the same number of visits but the highest income group has 23 per cent more visits. The number of visits for blacks falls midway between the low- and high-income whites.

For the age group 45–64, blacks experience 14 per cent fewer visits than whites and there are no income differentials for either race. The elderly experience an increasing number of visits as income increases, with the poor having a 25 per cent deficit relative to the highest income group. Blacks make 15 per cent fewer visits than whites (although this effect is just short of being statistically significant at the 5 per cent level).

The smoothed utilization patterns for those reporting good or excellent health (Table 5) are different in certain respects from the patterns for those reporting fair or poor health. For children there are similarities in that blacks experience fewer visits than whites (by 24 per cent) and the income effect is large (the poor have 29 per cent fewer visits than the nonpoor). The major differences are that the two lowest income groups exhibit the same number of visits and the remaining differences are smaller than those for children in fair or poor health.

For the 17–44 and 45–64 age groups there are no significant differences by race or income. For the elderly there are no differences by income but blacks have 18 per cent more visits than whites.

To summarize, the data show that, before adjustment for health status, blacks and the poor aged 17–64 use more physician visits than their counterparts. The effect of stratifying by self-assessed health status is to reverse or at least eliminate these differences. For children even the unadjusted data showed large income and racial differentials, but these gaps increase for those reporting fair or poor health. Furthermore, the elderly who report fair or poor health exhibit a steady increase in visits with increasing income, and blacks in this group also appear to have fewer visits. Although the stratification of the population into two health status categories (fair or poor versus good or excellent) does not result in groups with equal need for medical care, even such a crude adjustment suggests income and racial differences that are not apparent in the unadjusted data.

TABLE 5. Estimated* Number of Physician Visits per Person per Year for Those Reporting Good or Excellent Health by Age, Race and Income, 1976–78

Race and Income	Age Adjusted†	Under 17	17–44	45–64	65 and Older
White					
Below poverty	3.99	3.27	4.09	4.16	5.21
100–150% poverty	3.99	3.27	4.09	4.16	5.21
150–200% poverty	4.13	3.75	4.09	4.16	5.21
Above 200% poverty	4.37	4.60	4.09	4.16	5.21
Black					
Below poverty	3.87	2.49	4.09	4.16	6.17
100–150% poverty	3.87	2.49	4.09	4.16	6.17
150–200% poverty	3.98	2.86	4.09	4.16	6.17
Above 200% poverty	4.16	3.51	4.09	4.16	6.17

* Estimated from generalized log-linear model.
† Age-adjusted by the direct method.

Use–Bed Disability Ratio

The age-adjusted number of physician visits per 100 bed-disability days (UBDR) is 80 per cent higher among high-income whites than among those below the poverty level. For blacks, the differential is about 40 per cent. However, there are significant differences by age and race. Among children, whites have a higher UBDR than blacks in all income groups except those below poverty. Furthermore white children show a sharp increase in UBDR as income increases: the UBDR for the highest income group is nearly twice that for the poor. For black children, however, there is no evidence of an income differential.

In the 17–44 age group the UBDR for whites is 26 per cent higher than that for blacks in each income group. Both whites and blacks show a strong income effect: the UBDR for the high income group is 56 per cent higher than that for the poor.

White persons aged 45–64 generally have a higher UBDR than blacks, although the difference varies with income. Whites show a very strong increase in UBDR with increasing income (the ratio of the UBDR for the highest to lowest income groups is 2.5). The income effect for blacks appears to be limited to a deficit of 34 per cent among the poor, with no difference among the three higher income groups.

Blacks and whites in the oldest age group also have different trends by income. Among whites the UBDR shows a steady increase with a high-to-low ratio of 1.76. For blacks, however, there appears to be no difference within the two highest and two lowest income groups. The ratio between these high and low values is 2.18.

It may be useful to note at this point that these results differ from the CHAS–NORC data, which show no income differentials in the use-disability ratio.[1] The reasons for this discrepancy are probably related to the different survey methodologies previously discussed: the greater recall bias and omission of telephone contacts in CHAS–NORC and the use of broader income classes with no adjustment for family size by Aday et al.[1]

Sources of Care

The quantity of physician services used reflects only a part of the concern over access to care. As David Rogers remarked in his testimony to the Subcommittee on Health and Scientific Research of the United States Senate:

> Simply counting the number of visits people make to doctors each year is not

TABLE 6. Estimated* Number of Physician Visits per 100 Bed-Disability Days by Age, Race and Income, 1976–78

Race and Income	Age Adjusted†	Under 17	17–44	45–64	65 and Older
White					
Below poverty	53.0	53.5	64.5	38.4	34.9
100–150% poverty	63.5	69.9	73.8	44.2	44.2
150–200% poverty	73.9	82.1	85.6	51.2	50.8
Above 200% poverty	95.1	99.8	100.7	94.5	61.5
Black					
Below poverty	48.5	63.3	51.3	32.2	29.4
100–150% poverty	54.9	63.3	58.7	48.6	29.4
150–200% poverty	62.3	63.3	68.0	48.6	64.0
About 200% poverty	67.2	63.3	80.1	48.6	64.0

* Estimated from generalized log-linear model.
† Age-adjusted by the direct method.

sufficient. The term "physician visit" can mean many things. It can be a hurried visit to a "Medicaid mill" in which a patient is ping-ponged among a series of doctors, nurses and others to maximize reimbursement income. It can be hours spent waiting in a crowded, noisy hospital outpatient clinic or a few minutes with a doctor you've never seen before. At the other end of the scale–and far more desirable personally and medically–is the kind of visit you and I are accustomed to: seeing a doctor we know by name and who knows us and our families. Implicit in this relationship is the physician's assumption of personal responsibility for attending to an individual's medical care needs.[15]

The aggregate visit data discussed earlier encompass wide variation by source of care. Table 7 shows certain characteristics of visits by race and income (observed data are shown since the large and systematic differences in the observed data obviated the need to use the smoothing techniques discussed previously). Although the majority of visits take place in an office setting regardless of race or income, blacks are twice as likely as whites to use hospital outpatient departments or emergency rooms and half as likely to use the telephone. Furthermore, within each race, there is a consistent trend by income. These differentials are particularly large for children, especially those under age 6. In this group 25 per cent of the visits for whites above twice the poverty level are by telephone, compared with 14 per cent of those below the poverty level. The corresponding proportions for blacks are 20 per cent and 4 per cent. The range in proportion of visits at hospital clinics was 7 per cent (high income) to 18 per cent (poverty) for whites and 16 per cent to 36 per cent for blacks.

Conclusions

The NHIS data for 1976-78 indicate that income and race continue to influence the use of ambulatory medical care. Without adjusting for health status, the 1976-78 NHIS shows that the poor have more physician visits than those with higher incomes. After adjusting for age and health status, however, these differences are reversed. Depending on which measure is used, the poor have between 7 per cent and 44 per cent fewer visits than those with income above twice poverty level. Furthermore, the age- and health-adjusted data show that blacks have significantly fewer visits than their white counterparts.

For children under 17, even the unadjusted data show a sizeable deficit in

TABLE 7. Characteristics of Physician Visits by Race and Income, 1976–78

Race and Income	Physician's Office (% of visits)	Telephone (% of visits)	Hospital Clinic or Emergency Room (% of visits)
White			
Below poverty	64.1	10.8	15.3
100–150% poverty	67.2	11.1	14.0
150–200% poverty	68.7	12.9	13.3
Above 200% poverty	71.0	13.6	9.8
Black			
Below poverty	52.7	3.6	31.5
100–150% poverty	56.4	5.8	26.7
150–200% poverty	60.1	6.6	22.8
Above 200% poverty	64.8	7.9	18.7

physician visits among blacks and the poor. After adjusting for health status the income and racial differentials become larger. The unadjusted data for the elderly show no income or race effects. For the elderly who report fair or poor health, however, deficits among the poor and blacks become evident.

The methods used here to adjust utilization for health status are admittedly rather crude. Nonetheless, compared with the unadjusted data, they resulted in a different perspective on whether the quantity of ambulatory care received by the poor is commensurate with their need. Furthermore, these conclusions are consistent with the results of analyses using other methods to adjust utilization data through multiple regression and its econometric variants. These methods are generally used to simultaneously adjust for a large number of variables, including health status. Several studies using these methods have also shown that utilization of physician services by the poor lags substantially behind that of the nonpoor.[2]

Despite the plethora of studies on the measurement of use relative to need, the state of the art is still in flux. No single method or measure has achieved universal acceptance. The difference in results between the apparently similar use-disability ratio (CHAS–NORC, 1976) and use-bed disability ratio (NHIS, 1976–78) underscores the danger in relying on a single methodology in addressing this issue.

One set of differences between the poor and nonpoor that holds up consistently across many different studies and over several years are those involving characteristics of care. In all age groups there are substantial race and income differences in the type of visits. According to the 1976-78 NHIS data, blacks and the poor are more likely to use hospital clinics and less likely to use private physician offices or telephone consultations. The 1976 CHAS–NORC survey documented substantial deficits by the poor in access to regular source of care and satisfaction with medical care.[1,2]

The second consistent result is a deficit by the poor in use of preventive procedures. Several studies have shown that, compared to those with higher incomes, the poor are less likely to receive breast exams, Pap tests, prenatal care or immunization against childhood diseases (particularly in the preschool years).[2] Dental care is another area where large differences remain and for which little progress has been made over the past 15 years.[2,16,17]

These indications that substantial disparities in utilization and access to care still exist for low-income as well as other disadvantaged groups raise significant policy concerns. Aday et al.[1] suggest the following definition: "an equitable distribution of health care services is one in which illness is the major determinant of the allocation of resources." This does not imply that medical care alone can cure all illness or alleviate all pain. Nutrition, environmental exposures, housing, personal habits, education and other factors influence health status. Even if medical care were allocated in an equitable fashion, equal health outcomes for the poor and nonpoor could not be assured.

Nevertheless, medical care can contribute to the reduction of certain types of mortality, morbidity and functional impairment. Medical care can also alleviate pain and anxiety, perhaps explaining why public opinion polls have shown that a large proportion of the U.S. population places a high value on medical care. It would appear from the present evidence that still further progress is required to achieve the goal of equity in the distribution of medical care services.

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Appendix

In this section we present the details of the model-fitting procedures used to obtain the smoothed estimates given in the body of the article. A complete discussion of the rationale for applying linear model techniques to survey data can be found in references 12–14. Freeman et al.[13] point out the need to use appropriate estimates of the complete covariance matrix in such analyses. Unfortunately the computations involved in estimating such matrices were beyond the resources available for this study. Instead we use approximate variance curves derived for HIS[6] and assume zero covariance between cell estimates. Freeman et al.[13] found that the zero-covariance assumption tends to result in a loss of power. Thus, the test statistics we use are generally conservative and we run the risk of omitting significant effects. On the other hand, the multiple significance tests being carried out on the data may result in detecting too many "significant" differences.

We use a backward elimination procedure in model fitting by starting with a fully saturated model and reducing interaction terms that are not significant (at the 10 per cent level) in a hierarchical manner (i.e., three-way interactions, then two-way, etc.). If interactions with age were present we fit models separately within age group. Although this is not a standard practice in formal statistical modeling, we felt the nature of the measures used reflected substantive differences by age (e.g., the type of

poor health reported for a child may differ from that reported for an elderly person). If fitted income effects did not follow a monotonic pattern, we tested adjacent categories to see if they could be combined.

After eliminating all nonsignificant terms we examined the table of residuals to look for unusual patterns or outliers. Since the variances of estimates for blacks were substantially larger than those for whites, residuals for blacks were generally larger. If the examination of residuals suggested a pattern, terms were added to the model to improve its fit. While this approach invalidates a strict statistical-significance-level interpretation of the effects, we feel it provides a reasonable compromise between formal modeling and minimizing the risk of omitting real differences.

Table A1 shows the observed means or proportions for each cell in the age × race × income tables and Table A2 shows the corresponding standard errors. The results of the model-fitting procedures for each variable are outlined in tables A3–A7 and discussed in the following text. The abbreviations used in the tables are as follows:

P	=	below poverty level
1.5P	=	100—150 per cent poverty
2P	=	150—200 per cent poverty
NP	=	more than 200 per cent poverty
POP	=	population (in thousands)
%FP	=	per cent reporting fair or poor health
MDV	=	number of physician visits
% no MDV	=	per cent with no physician visits in 2 years preceding interview
MDV/FP	=	number of physician visits for those reporting fair or poor health
MDV/GE	=	number of physician visits for those reporting good or excellent health
MDV/BD	=	number of physician visits per 100 bed-disability days
A	=	age
R	=	race
I	=	income

Self-Assessed Health

Table A3 shows the results of model fitting for the proportion reporting fair or poor health. After eliminating the three-way interaction terms, each of the two-way interactions were highly significant. Since the race × income interaction terms within each age were also significant, we did not attempt to reduce the model with no three-way interaction any further. The residuals were all below 5 per cent for whites and 14 per cent for blacks.

Physician Visits

Main effects models within each age group provided an adequate fit to the data on number of physician visits (Table A4). For children under 17 and adults age 17–44, two adjacent income categories could be combined. Residuals were all below 3 per cent for whites and, except for two cases, below 10 per cent for blacks.

For the per cent with no physician visit in the past two years (Table A5), the combined race × income effect, given no three-way interaction, was significant at the 0.066 level. However, there was no evidence of a race × income interaction within any age group, so we proceeded to fit separate main effects models within each age. The residuals were below 3 per cent for whites. Among blacks there were three residuals between 15 per cent and 20 per cent, but all the rest were below 10 per cent.

For those reporting fair or poor health, the model with ARI = RI = 0 fit well (Table A6). However, when main-effects models were fit separately within each group, there was a significant race × income interaction for those 17–44. This interaction resulted from a positive relationship between visits and income for whites but a negative relationship for blacks in the 17–44 age group. When these income effects are examined separately, the effect for whites is significant ($p = 0.002$), while that

TABLE A1. Selected Health Characteristics by Age, Race and Income, 1976–78

	POP	% FP	MDV	% No MDV	MDV/FP	MDV/GE	MDV/BD
Both races							
All ages							
P	29560.36	22.06	5.07	14.71	9.75	3.74	46.82
1.5P	23311.58	17.73	4.95	14.74	10.19	3.82	54.80
2P	38203.75	12.30	4.73	13.08	10.37	3.94	67.28
NP	98763.69	7.59	4.93	11.40	12.09	4.34	92.31
Under 17							
P	11237.52	8.10	3.42	14.61	7.66	3.05	55.44
1.5P	7222.80	4.96	3.40	13.51	10.18	3.04	66.63
2P	12390.35	3.43	3.97	10.51	12.59	3.66	81.68
NP	23023.05	2.61	4.85	6.99	16.50	4.53	98.09
17–44							
P	10561.60	18.56	5.[illegible]	13.66	10.98	4.12	59.73
1.5P	8131.35	11.55	4.[illegible]2	14.54	11.02	4.01	70.71
2P	15187.08	8.71	4.46	13.45	10.27	3.90	83.16
NP	45042.41	5.09	4.57	12.05	12.26	4.16	99.12
45–64							
P	3680.25	50.98	7.39	17.88	10.54	4.12	36.62
1.5P	3374.80	37.52	6.25	18.55	10.51	3.69	45.44
2P	6033.45	27.58	5.53	17.21	9.77	3.92	49.64
NP	24984.88	13.74	5.22	14.22	11.29	4.25	80.91
65 and over							
P	4080.99	43.46	6.66	14.87	8.62	5.15	33.57
1.5P	4582.63	34.25	6.[illegible]5	14.24	9.45	5.19	42.49
2P	4592.87	28.00	6.62	13.36	10.52	5.10	50.98
NP	5713.43	20.35	6.80	11.71	11.80	5.52	62.26
White							
All ages							
P	20759.48	21.76	5.21	15.39	9.93	3.90	47.67
1.5P	19346.57	17.15	4.99	14.83	10.33	3.[illegible]8	55.58
2P	34548.60	11.88	4.73	13.23	10.48	3.95	68.04
NP	92787.44	7.34	4.95	11.55	12.36	4.37	94.44
Under 17							
P	7188.61	7.30	3.74	15.00	8.86	3.33	53.53
1.5P	5962.82	4.33	3.56	13.30	12.00	3.17	69.85
2P	11185.88	3.27	4.02	10.46	13.24	3.71	82.07
NP	21569.42	2.50	4.94	7.08	17.19	4.62	99.76
17–44							
P	7600.13	16.73	5.15	14.52	10.18	4.14	64.00
1.5P	6890.39	10.56	4.69	14.67	10.73	3.98	72.04
2P	13612.24	8.22	4.41	13.73	10.31	3.88	84.71
NP	42009.13	4.83	4.57	12.22	12.69	4.16	102.03
45–64							
P	2674.02	49.62	7.63	19.13	11.18	4.13	38.35
1.5P	2815.03	36.34	6.21	19.29	10.46	3.79	44.19
2P	5416.59	26.61	5.54	17.50	9.82	3.99	51.18
NP	23692.88	13.25	5.23	14.41	11.51	4.27	94.47
65 and over							
P	3296.72	42.27	6.59	15.20	8.91	4.89	34.88
1.5P	4178.33	33.37	6.70	14.27	9.73	5.18	44.19
2P	4333.89	27.19	6.53	13.43	10.58	5.02	50.79
NP	5516.08	19.97	6.76	11.65	11.82	5.50	61.46
Black							
All ages							
P	8800.87	22.77	4.74	13.13	9.34	3.38	44.75
1.5P	3465.01	21.07	4.71	14.25	9.54	3.43	50.52

TABLE A1. Continued

	POP	% FP	MDV	% No MDV	MDV/FP	MDV/GE	MDV/BD
2P	3655.15	16.24	4.74	11.70	9.65	3.79	60.91
NP	5976.26	11.41	4.56	9.00	9.32	3.94	66.85
Under 17							
P	4048.92	9.52	2.87	13.93	6.03	2.54	60.42
1.5P	1259.98	7.92	2.64	14.46	5.45	2.40	51.49
2P	1204.47	4.89	3.46	10.98	8.61	3.20	77.72
NP	1453.64	4.24	3.46	5.73	10.44	3.15	72.40
17–44							
P	2961.47	23.27	6.02	11.44	12.45	4.07	52.11
1.5P	1240.96	17.06	5.56	13.87	12.02	4.23	65.08
2P	1574.84	12.98	4.88	11.05	10.07	4.11	75.75
NP	3033.28	8.74	4.65	9.57	8.96	4.24	71.42
45–64							
P	1006.22	54.60	6.76	14.54	8.99	4.07	32.24
1.5P	559.77	43.49	6.44	14.84	10.73	3.14	52.70
2P	616.86	36.10	5.46	14.59	9.41	3.23	39.14
NP	1292.00	22.73	5.06	10.65	8.95	3.92	53.10
65 and over							
P	784.27	48.49	6.92	13.50	7.54	6.35	29.10
1.5P	404.30	43.29	6.20	13.99	7.24	5.40	29.72
2P	258.98	41.48	8.10	12.11	9.92	6.80	53.73
NP	197.35	31.05	7.83	13.42	11.57	6.14	90.42

TABLE A2. Standard Errors of Estimates in Table A1

	POP	% FP	MDV	% No MDV	MDV/FP	MDV/GE	MDV/BD
Both races							
All ages							
P	109.65	0.25	0.08	0.21	0.24	0.07	1.18
1.5P	113.26	0.26	0.09	0.24	0.31	0.08	1.61
2P	89.16	0.18	0.07	0.18	0.30	0.06	1.74
NP	0.0	0.09	0.04	0.11	0.26	0.04	1.31
Under 17							
P	96.45	0.27	0.10	0.35	0.56	0.10	2.67
1.5P	81.52	0.27	0.12	0.42	1.06	0.12	4.23
2P	99.64	0.17	0.10	0.29	1.12	0.10	4.00
NP	113.23	0.11	0.08	0.18	1.13	0.08	3.50
17–44							
P	94.37	0.39	0.13	0.35	0.48	0.13	2.46
1.5p	85.50	0.37	0.14	0.41	0.69	0.14	3.66
2P	105.80	0.24	0.10	0.29	0.55	0.10	3.52
NP	47.89	0.11	0.06	0.16	0.47	0.06	2.73
45–64							
P	60.71	0.86	0.27	0.66	0.47	0.27	1.84
1.5P	58.34	0.87	0.26	0.70	0.58	0.24	2.75
2P	75.60	0.60	0.18	0.51	0.48	0.17	2.48
NP	113.08	0.23	0.09	0.23	0.37	0.08	2.95
65 and over							
P	63.63	0.81	0.24	0.58	0.43	0.28	1.65
1.5P	67.04	0.73	0.23	0.54	0.48	0.24	2.11
2P	67.10	0.69	0.23	0.52	0.37	0.23	2.68
NP	73.85	0.56	0.21	0.44	0.65	0.20	3.10

TABLE A2. Continued

	POP	% FP	MDV	% No MDV	MDV/FP	MDV/GE	MDV/BD
White							
All ages							
P	112.43	0.30	0.09	0.26	0.29	0.09	1.38
1.5P	111.81	0.28	0.09	0.26	0.35	0.09	1.75
2P	100.45	0.18	0.07	0.19	0.32	0.07	1.84
NP	0.0	0.09	0.04	0.11	0.28	0.04	1.39
Under 17							
P	81.36	0.32	0.13	0.44	0.80	0.13	3 03
1.5P	75.22	0.27	0.14	0.46	1.39	0.13	4.84
2P	96.29	0.18	0.11	0.30	1.24	0.11	4.19
NP	112.84	0.11	0.09	0.18	1.22	0.09	3.66
17–44							
P	83.22	0.45	0.15	0.42	0.56	0.15	3.21
1.5P	79.95	0.39	0.15	0.44	0.77	0.15	4.12
2P	102.59	0.25	0.10	0.31	0.60	0.10	3.82
NP	71.19	0.11	0.06	0.17	0.51	0.06	2.92
45–64							
P	52.34	1.01	0.32 0.79		0.58	0.32	2.24
1.5P	53.62	0.94	0.28	0.78	0.64	0.26	2.90
2P	72.17	0.63	0.19	0.54	0.52	0.18	2.71
NP	113.27	0.23	0.09	0.24	0.39	0.08	3.18
65 and over							
P	57.71	0.90	0.27	0.65	0.49	0.29	1.93
1.5P	64.31	0.76	0.24	0.56	0.52	0.25	2.32
2P	65.38	0.70	0.23	0.54	0.60	0.23	2.76
NP	72.74	0.56	0.21	0.45	0.67	0.21	3.11
Black							
All ages							
P	85.19	0.47	0.13	0.38	0.42	0.13	1.95
1.5P	59.05	0.72	0.21	0.62	0.71	0.20	3.57
2P	60.52	0.64	0.21	0.55	0.80	0.20	4.45
NP	75.30	0.43	0.16	0.39	0.73	0.16	4.04
Under 17							
P	63.41	0.48	0.15	0.57	0.74	0.15	5.35
1.5P	36.49	0.79	0.26	1.03	1.37	0.25	8.03
2P	35.70	0.65	0.30	0.94	2.35	0.30	12.48
NP	39.12	0.35	0.28	0.64	2.59	0.27	10.33
17–44							
P	54.90	0.81	0.27	0.61	0.87	0.24	3.56
1.5P	36.23	1.11	0.39	1.02	1.53	0.37	7.64
2P	40.66	0.88	0.32	0.82	1.39	0.31	8.42
NP	55.52	0.53	0.23	0.56	1.13	0.23	6.09
45–64							
P	32.70	1.64	0.49	1.16	0.79	0.54	3.08
1.5P	24.51	2.18	0.64	1.57	1.33	0.56	7.96
2P	25.71	2.02	0.55	1.48	1.28	0.51	5.58
Np	36.94	1.21	0.36	0.89	1.08	0.36	5.98
65 and over							
P	28.94	1.86	0.36	1.27	0.85	0.74	3.03
1.5P	20.86	2.57	0.73	1.80	1.22	0.90	4.55
2P	16.72	3.19	1.08	2.11	1.90	1.27	10.69
NP	14.61	3.43	1.21	2.53	2.78	1.26	25.20

TABLE A3. Summary of Model-Fitting Procedure: Per Cent Reporting Fair or Poor Health

Model	df	χ^2	p
Saturated model			
ARI = 0*	9	10.09	0.343
Model with ARI = 0† (Final model)	9	10.09	0.343
AR = 0*	3	31.12	<0.001
AI = 0*	9	248.65	<0.001
RI = 0*	3	69.85	<0.001
Main effects models for each age:			
Under 17†	3	7.73	0.052
17–44†	3	12.05	0.007
45–64†	3	46.44	<0.001
65 and over†	3	13.73	0.003

* Hypothesis test given specified model.
† These lines give the goodness-of-fit χ^2 for the specified model.

for blacks is not (p = 0.102). There were, however, no significant differences among the three lowest income groups for whites (p = 0.844). Overall, residuals for whites were below 5 per cent except for one cell with a residual of 10 per cent. There were two large residuals (25 per cent) among blacks but both fitted values were within

TABLE A4. Summary of Model-Fitting Procedure: Number of Physician Visits per Year

Model	df	χ^2	p
Saturated model			
ARI = 0*	9	8.12	0.522
Model with ARI = 0†	9	8.12	0.522
AR = 0*	3	57.48	<0.001
AI = 0*	9	198.95	<0.001
RI = 0*	3	3.65	0.302
Main effects model for each age:			
Under 17†	3	2.93	0.403
R = 0*	1	44.42	<0.001
I = 0*	3	109.41	<0.001
P = 1.5P*	1	1.39	0.238
17–44†	3	4.55	0.208
R = 0*	1	11.06	<0.001
I = 0*	3	31.13	<0.001
2P = NP*	1	1.30	0.254
45–64†	3	1.49	0.684
R = 0*	1	0.93	0.335
I = 0*	3	83.01	<0.001
65 and over†	3	2.80	0.424
R = 0*	1	1.18	0.277
I = 0*	3	0.61	0.895
Final model†	19	16.98	0.591

* Hypothesis test given specified model.
† These lines give the goodness-of-fit χ^2 for the specified model.

TABLE A5. Summary of Model-Fitting Procedure: Per Cent With No Physician Visits in Past 2 Years

Model	df	χ^2	p
Saturated model			
ARI = 0*	9	5.53	0.786
Model with ARI = 0†	9	5.53	0.786
AR = 0*	3	17.16	<0.001
AI = 0*	9	191.79	<0.001
RI = 0*	3	7.18	0.066
Main effects model for each age:			
Under 17†	3	6.08	0.108
R = 0*	1	0.92	0.337
I = 0*	3	487.62	<0.001
17–44†	3	4.25	0.235
R = 0*	1	36.21	<0.001
I = 0*	1	68.26	<0.001
P = 1.5P*	1	1.07	0.301
45–64†	3	0.81	0.848
R = 0*	1	29.40	<0.001
I = 0*	3	97.58	<0.001
P = 1.5P*	1	0.04	0.841
65 and over†	3	1.56	0.668
R = 0*	1	0.65	0.420
I = 0*	3	23.06	<0.001
Final model†	16	15.39	0.496

* Hypothesis test given specified model.
† These lines give the goodness-of-fit χ^2 for the specified model.

the 95 per cent confidence interval for the cell estimate. The remaining residuals were below 10 per cent.

For those reporting good or excellent health, the main effects models within each age group fit very well (Table A7). With only two exceptions, residuals among both whites and blacks were below 13 per cent. The two exceptions were the middle-income groups among blacks aged 45–64 which had residuals of 25 per cent. The fitted value for both cells were within the 95 per cent confidence interval of their corresponding estimate.

Use-Bed-Disability Ratio

The three-way interaction was highly significant (Table A8) so we proceeded to fit separate models within each age group. For children, there was a significant race × income interaction, with whites showing strong income effects but blacks showing none. A main effects model provided an adequate fit in the 17–44 group with both race and income main effects highly significant.

Among adults 45–64 there also was a significant interaction. Whites showed a strong and consistent income effect. Blacks, however, showed no difference among the three highest income groups but a significantly lower UBDR among the poor. In the oldest age group the interaction was significant at the 0.073 level. Once again whites experienced a strong and consistent income effect but blacks showed a significant difference only between the two highest and two lowest income groups.

Residuals among the estimates for whites were all below 3 per cent. Among blacks, however, some of the residuals were rather large, although no smoothed estimate fell outside the 95 per cent confidence interval based on the corresponding (unsmoothed) cell value.

TABLE A6. Summary of Model-Fitting Procedure: Number of Physician Visits per Year for Those reporting Fair or Poor Health

Model	df	χ^2	p
Saturated model			
ARI = 0*	9	12.32	0.196
Model with ARI = 0†	9	12.32	0.196
AR = 0*	3	18.09	<0.001
AI = 0*	9	33.05	<0.001
RI = 0*	3	5.97	0.113
Main effects model for each age:			
Under 17†	3	1.67	0.642
R = 0*	1	18.81	<0.001
I = 0*	3	37.66	<0.001
17–44†	3	12.21	0.007
45–64†	3	3.27	0.353
R = 0*	1	5.46	0.019
I = 0*	3	5.52	0.138
65 and over†	3	1.14	0.768
R = 0*	1	3.52	0.061
I = 0*	3	16.93	<0.001
Saturated model for age 17–44			
RI = 0*	3	12.21	0.007
Whites			
I = 0*	3	14.66	0.002
P = 1.5P = 2P*	2	0.34	0.844
Blacks			
I = 0*	3	6.20	0.102
Final model†	17	18.13	0.381

* Hypothesis test given specified model.
† These lines give the goodness-of-fit χ^2 for the specified model.

TABLE A7. Summary of Model-Fitting Procedure: Number of Physician Visits per Year for Those Reporting Good or Excellent Health

Model	df	χ^2	p
Saturated model			
ARI = 0*	9	5.03	0.832
Model with ARI = 0†	9	5.03	0.832
AR = 0*	3	39.60	<0.001
AI = 0*	9	64.80	<0.001
RI = 0*	3	1.39	0.707
Main effects model for each age:			
Under 17†	3	3.27	0.352
R = 0*	1	39.97	<0.001
I = 0*	3	125.48	<0.001
P = 1.5P*	1	1.00	0.317
17–44†	3	.66	0.883
R = 0*	1	.44	0.507
I = 0*	3	5.64	0.131
45–64†	3	1.00	0.802
R = 0*	1	2.48	0.115
I = 0*	3	5.65	0.130
65 and over†	3	1.50	0.682
R = 0*	1	5.24	0.022
I = 0*	3	3.32	0.345
Final model†	24	25.56	0.376

* Hypothesis test given specified model.
† These lines give the goodness-of-fit χ^2 for the specified model.

TABLE A8. Summary of Model-Fitting Procedure: Use-Bed-Disability Ratio

Model	df	χ^2	p
Saturated model			
ARI = 0*	9	22.98	0.006
Separate saturated models within each age group:			
Under 17			
RI = 0*	3	8.10	0.044
Whites			
I = 0*	3	90.42	<0.001
Blacks			
I = 0*	3	4.55	0.208
17–44			
RI = 0*	3	3.43	0.330
Model with RI = 0†	3	3.43	0.330
R = 0*	1	20.30	<0.001
I = 0*	3	87.68	<0.001
45–64			
RI = 0*	3	14.84	0.002
Whites			
I = 0*	3	259.53	<0.001
Blacks			
I = 0*	3	14.45	0.002
1.5P = 2P = NP*	2	3.21	0.201
65 and over			
RI = 0*	3	6.96	0.073
Whites			
I = 0*	3	60.53	<0.001
Blacks			
I = 0*	3	20.43	<0.001
P = 1.5P*	1	0.01	0.920
2P = NP*	1	2.31	0.129
Final model†	10	13.51	0.197

* Hypothesis test given specified model.
† These lines give the goodness-of-fit χ^2 for the specified model.

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Improving Hospital Discharge Data: Lessons From the National Hospital Discharge Survey

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The most recent study by the Institute of Medicine of the reliability of hospital discharge data abstracted from patients' medical records documents the continuing presence of a substantial level of imprecision and error, affirming the findings of earlier studies. The most obvious factors leading to unreliability stem from inadequacies in the face sheets of medical records. Significant improvements could be attained if the discharge summary were routinely used to abstract information on patient disposition and principal diagnosis and if the operative report were routinely used to abstract information on principal procedure. Additional recommendations are offered for developmental activities intended to improve the designation of diagnosis, diagnostic classification schemes and hospital medical records systems.

SEVERAL RECENT STUDIES have questioned the reliability of hospital discharge data and, in particular, information on diagnoses and procedures.[1–6] The latest in a series of studies by the Institute of Medicine examined data collected by the National Hospital Discharge Survey (NHDS), which is conducted by the National Center for Health Statistics (NCHS) and constitutes the only source of nationally aggregated hospital discharge information for the total noninstitutionalized population. The scope of this last study was broadened to permit greater attention to methods for data improvement.

The NHDS is a voluntary survey that has been conducted since 1964 and yields statistics on the utilization of all general and special, nonfederal, short-stay hospitals in the United States and on characteristics of patients who use those services. The information is abstracted primarily from the face sheets of patients' medical records by employees of participating hospitals or by the Bureau of the Census, which acts as an agent for NCHS. The abstracted data are forwarded to NCHS facilities in North Carolina, where they are centrally coded and processed.[7–9]

The division of responsibility for abstracting and coding influences the qual-

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ity of the data, because NCHS coders do not have the opportunity to consult the medical record and instead base their code on the handwritten information recorded on the abstract by the person responsible for abstracting in the hospital. The process is further complicated by the inherent complexities of diagnostic information and the wide variation among hospitals in the components, organization and quality of the medical record. This study was designed to take these factors into consideration.

Study Methods

The methods for this study were similar to those of the previous IOM studies, which have been reported elsewhere.[10] Very briefly, the 66 participating hospitals were selected by sampling the NHDS master sample. Within each participating hospital a diagnostic-specific probability sample of NHDS abstracts from calendar year 1977 was drawn, which could be linked to the original medical record within that hospital. The sampling plan was designed to facilitate comparison with some specific diagnoses included in earlier IOM studies and also permitted the aggregation of diagnoses into broader categories so that the reliability of published NCHS statistics could be examined.

The field work was conducted by specially trained Registered Record Administrators (RRAs), who independently reabstracted all patient records selected within participating hospitals, relying primarily on the Uniform Hospital Discharge Data Set definitions. Results of the reabstracting were compared with information on a computerized version of the NHDS abstract, which is part of the NHDS file maintained by NCHS. Discrepancies between the two abstracts were noted. The medical records were then re-examined, along with both the computerized and original hard-copy NHDS abstracts, in an attempt to understand the reasons for discrepancies.

The original study methods were modified slightly to better address the unique features of the NHDS and to build on earlier findings. More emphasis was placed on identifying the location within the medical record of the correct information and, particularly, on assessing the adequacy of the face sheet for abstracting discharge information. More detailed analyses were performed of information on principal procedures. As in previous studies, the diagnostic-specific sampling plan included the total range of the ICDA-8 classification scheme; however, the NHDS study approximated a self-weighting sample more closely than previously.

Nonmedical Data Reliability

As shown in Table 1, the NHDS information was highly reliable for dates of hospital admission and discharge, patient's age or date of birth, sex, race, marital status and patient disposition at the time of hospital discharge. Similar levels of reliability were obtained for most nonmedical data items in previous studies‡ (Table 2). The somewhat lower figures for NHDS data on payment source reflect ambiguity in the procedures for distinguishing between expected principal and additional sources of payment; this has since been alleviated by a routine NCHS computer edit.

Most NHDS discrepancies on dates of hospital admission and discharge, date of birth or age, and sex stemmed from clerical problems, such as transposing code numbers or erroneously recording the year in which the abstracting occurred in place of the year of admission. Most discrepancies

‡ The 1977 IOM study of discharge abstracts examined abstracts processed by four private abstract services, reflecting care provided to a national sample of Medicare and Medicaid patients during 1974. The 1977 IOM study of medicare records examined data on utilization by a national sample of Medicare beneficiaries discharged during 1974. The records were generated by the Health Care Finance Administration from claims information submitted by participating hospitals to receive Medicare reimbursement.

TABLE 1. Discrepancy Between NHDS Abstract and IOM Reabstract and the Correct Data Source for Nonmedical Data (Weighted Per Cent)

Data Item	No Discrepancy	Correct Data Source if Discrepant				
		NHDS Abstract	IOM Re-Abstract	Either	Neither	Total
Admission date	99.3	0.4	0.2	0.1	—	100.0
Discharge date	97.9	1.6	0.5	—	—	100.0
Date of birth or age	96.4	1.4	1.7	0.5	—	100.0
Sex	99.5	0.3	0.2	—	—	100.0
Race	92.5	1.1	3.0	3.4	—	100.0
Marital status	95.5	1.7	1.5	1.3	—	100.0
Disposition of patient	92.9	0.4	6.2	0.5	—	100.0

Unweighted N = 3313.

associated with race, marital status and patient disposition were traced to inadequacies of the face sheet of the medical record.

Analyses of Diagnostic Information

Reliability of the NHDS diagnostic information varied considerably, depending on the diagnostic classification and coding schemes that formed the basis of comparison. In Table 3, diagnostic data from all three studies are presented. Four-digit comparisons are quite similar across studies; reliability increases as more inclusive diagnostic groupings are used.

The reliability of individual diagnoses also varied by diagnostic classification and coding. Table 4 shows diagnostic-specific data from all three studies and gives the reliability figures for the most specific, as well as the most inclusive diagnostic groupings within each study. In all studies chronic ischemic heart disease is the most troublesome specific diagnosis. Again, levels of reliability increase as more inclusive diagnostic groupings are used.

TABLE 2. Discrepancies Between the Institute of Medicine Field Team and Original Data Source (Private Abstract, Medicare Record or National Hospital Discharge Survey Abstract) for Nonmedical Data

Data Item	Weighted Per Cent with No Discrepancy		
	NHDS Abstract	Medicare Record*	Private Abstract†
Admission Date	99.3	99.5	99.7
Discharge date	97.9	99.3	99.2
Date of birth or age	96.4	—	97.7
Sex	99.5	99.4	99.1
Payment source	75.6–87.0	—	98.1
Race	92.5	—	—
Marital status	95.5	—	—
Disposition of patient	92.9	—	—

* Reflects claims information for Medicare discharges during 1974.
† Reflects abstract information for Medicare and Medicaid discharges during 1974.

The presence of multiple diagnoses in a single patient influences the reliability of data. Approximately 55 per cent of the cases in the NHDS sample had multiple diagnoses. For those cases, the reliability of the principal diagnosis was 54.7 per cent. This contrasts with a reliability of 73.5 per cent for the cases that did not have additional diagnoses and confirms the findings in the Medicare study. Diagnostic categories whose reliability was notably lowered with multiple diagnoses included diseases of the nervous system and sense organs, diseases of the circulatory system, complications of pregnancy, childbirth and the puerperium and symptoms and ill-defined conditions.

For some cases, particularly when the patient had many medical problems, it was impossible for the field team to determine which diagnosis should be regarded as principal and which data source (the IOM or the original data) should be viewed as correct. For some of these patients, each diagnosis may have contributed equally to the necessity for hospital admission, so that the designation of a single, principal diagnosis would have been artificial and possibly misleading. The per cent of these cases varied, again depending on the level of diagnostic coding refinement. For all records examined in the NHDS study, the figures ranged from 6.5 per cent with four-digit comparisons to 2.0 per cent with the broader NCHS classes. This compares with 4.6 per cent for Medicare data and 10.7 per cent for the private services, both using four-digit coding.

Some uses of hospital discharge data depend on the correct inclusion of a specific diagnosis among the list of diagnoses appearing on the NHDS abstract, regardless of the designation of a principal diagnosis. For almost every diagnosis examined, the likelihood that it would be appropriately included on the NHDS abstract was greater than the likelihood of it being correctly designated as principal. Nevertheless, the more troublesome diagnoses in each analysis tended to be the same—particularly symptoms and ill-defined conditions (Table 5).

TABLE 3. Comparisons Between IOM and Original Data Source for All Diagnoses Combined by Level of Coding Refinement

	Weighted Per Cent with No Discrepancy		
	NHDS Data	Medicare Record*	Private Abstract*
Four-Digit	63.4	58.4	65.2
Three-Digit	73.2	61.7	74.0
CPHA List A	—	—	70.2
DRGs	—	80.8	—
NCHS Categories	86.0	—	—

* See Table 2.

The influence of diagnostic reliability on utilization statistics was estimated by calculating diagnostic-specific hospital discharge rates. Using the broad diagnostic classes appearing in the NCHS Series 13 publications, the NHDS data overestimate discharge rates for slightly more than half the classes in amounts generally ranging from about 3 to 37 discharges per 10,000 population. The largest difference occurs for diseases of the circulatory system, which is caused primarily by the overcoding of chronic ischemic heart disease. Conversely, for slightly under half the diagnostic classes the NHDS data underestimate discharge rates in amounts ranging from 2 to 40 discharges per 10,000 population. The largest undercount occurs for patients with symptoms and ill-defined conditions. Based on IOM findings, the NHDS data for these conditions should have been augmented by 40 discharges per 10,000 population. Next in importance were patients with special conditions and examinations without sickness (or tests with negative findings), for whom the NHDS data should have been increased by about 17 discharges per 10,000 population.

TABLE 4. Comparisons Between the IOM Field Team and Original Data Source by Diagnosis at Differing Levels of Coding Refinement

NCHS Diagnostic Category	Weighted Per Cent with No Discrepancy*					
	NHDS Data		Medicare Record		Private Abstract	
	Four Digit	NCHS Groups	Four Digit	DRGs	Four Digit	CPHA List A
Infective and Parasitic Diseases	60.8	79.4	—	—	—	—
Neoplasms	52.4	74.3	—	—	—	—
Malignant neoplasm of bronchus and lung	—		79.9	79.9	—	—
Carcinoma of the breast	—		—	—	75.7	79.0
Endocrine, Nutritional and Metabolic Diseases	26.3	67.7	—	—	—	—
Diabetes	10.7†	72.1	49.7	56.2	60.8	61.9
Diseases of the Blood and Blood Forming Organs	66.8	72.6	—	—	—	—
Mental disorders	61.7	86.2	—	—	—	—
Neuroses	—		—	—	80.1	80.5
Diseases of the Nervous System and Sense Organs	68.4	79.8	—	—	—	—
Cataract	—		97.4	97.9	94.3	94.2
Diseases of the Circulatory System	53.1	79.7	—	—	—	—
Cerebrovascular accident	43.2	73.6	58.5	84.7	78.4	84.0
Congestive heart failure	68.8	85.4	58.5	60.6	—	—
Chronic ischemic heart disease	37.5	81.4	36.8	38.6	30.2	33.6
Acute myocardial infarction	—	—	67.3	76.1	68.2	93.9
Diseases of the Respiratory System	72.1	92.4	—	—	—	—
Bronchitis	72.2	95.7	89.8	95.7	—	—
Bronchopneumonia	79.6	93.2	75.9	80.5	—	—
Diseases of the Digestive System	70.5	86.2	—	—	—	—
Cholelithiasis/cholecystitis	78.4	94.8	62.8	84.4	73.2	93.8
Intestinal obstruction without mention of hernia	—	—	58.1	68.8	—	—
Diverticulosis of intestine	—	—	86.5	86.9	—	—
Diseases of the Genitourinary System	74.7	87.4	—	—	—	—
Inguinal hernia without mention of obstruction	—	—	96.7	96.7	89.1	92.7
Hyperplasia of the prostate	—	—	87.1	87.1	—	—
End stage renal disease	—	—	—	—	64.9	66.4
Complications of Pregnancy, Childbirth and the Puerperium	64.3	98.0	—	—	—	—
Diseases of Skin and Subcutaneous Tissue	67.9	77.5	—	—	—	—
Diseases of the Musculoskeletal System and Connective Tissue	64.3	86.3	—	—	—	—
Fracture, neck of femur	—	—	70.6	93.4	81.1	97.6
Displacement of intervetebral disc			—	—	77.8	70.9
Congenital Anomalies	49.5	59.8	—	—	—	—
Certain Causes of Perinatal Morbidity and Mortality	74.3	85.3	—	—	—	—
Symptoms and Ill-defined Conditions	54.6	60.5	—	—	—	—
Hypertrophy of tonsils and adenoids	—	—	—	—	97.0	97.0
Accidents, Poisonings and Violence	61.7	89.6	—	—	—	—
Special Conditions and Examinations Without Sickness (or tests with negative findings)	72.8	79.6	—	—	—	—
Newborns	63.0	96.0	—	—	—	—
Deliveries	—	—	—	—	86.5	89.4

* Blanks indicate that the data in the original study were not analyzed in a manner that would permit completing the cell; they do *not* indicate that the relevant diagnoses were omitted from the sample, although in some instances the unweighted n's may have been quite small. All three studies were designed to reflect the total range of diagnoses included in ICDA-8, which was essential to compute net and gross difference rates. The NHDS study placed more emphasis on the broader diagnostic categories than specific diagnosis, whereas the two other studies focused more on the latter.

† This per cent is particularly low because the fourth digit of the NHDS modification is inconsistent with ICDA-8, which was used by the IOM field team.

TABLE 5. Frequency of the IOM Principal Diagnosis Appearing Somewhere on the NHDS Computerized Abstract vs. Frequency of Agreement on Principal Diagnosis by Three-Digit Comparison (Weighted Per Cent)*

IOM Principal Diagnosis	IOM Principal Diagnosis	
	Appears on NHDS Abstract	Agrees with NHDS Principal
Infective and Parasitic Diseases	68.5	62.6
Neoplasms	70.5	57.4
Endocrine, Nutritional and Metabolic Diseases	97.0	71.3
Diabetes	99.8	65.7
All other endocrine, nutritional and metabolic diseases	90.6	84.2
Diseases of the Blood and Blood Forming Organs	92.3	58.2
Mental Disorders	90.3	80.8
Diseases of the Nervous System and Sense Organs	90.0	85.7
Diseases of the Circulatory System	82.5	74.3
Cerebrovascular accident	72.5	61.2
Congestive heart failure	87.7	73.9
Chronic ischemic heart disease	95.1	93.3
All other circulatory diseases	81.3	66.6
Diseases of the Respiratory System	81.6	73.6
Bronchitis	75.3	66.4
Bronchopneumonia	79.4	76.1
All other respiratory diseases	86.4	75.3
Diseases of the Digestive System	85.8	81.0
Cholelithiasis/cholecystitis	90.5	85.8
All other digestive diseases	84.7	79.9
Diseases of the Genitourinary System	87.6	78.1
Complications of Pregnancy, Childbirth and the Puerperium	76.4	75.4
Diseases of the Skin and Subcutaneous Tissue	87.5	73.0
Diseases of the Musculoskeletal System and Connective Tissue	78.4	71.3
Congenital Anomalies	80.0	80.0
Certain Causes of Perinatal Morbidity and Mortality	48.2	20.5
Symptoms and Ill-Defined Conditions	30.4	29.7
Accidents, Posionings and Violence	73.8	67.5
Special Conditions and Examinations without sickness (or tests with negative findings)	50.1	48.6
Newborns	99.2	99.2
Total	80.2	73.1

* In interpreting this table it is important to note that cases were grouped into diagnostic categories according to the diagnosis designated as principal on the IOM reabstract. In previous tables, the categorization is based on NHDS diagnoses. This distinction is important in interpreting diagnostic-specific discrepancy rates for NHDS data, which should be taken from Table 4 and not from Table 5.

Similar comparisons were made for average-length-of-stay (ALOS). For seven of the 19 NCHS diagnostic classes, the NHDS data resulted in ALOS figures that were larger than the IOM by amounts ranging from 0.1 to 1.4 days—the latter was for patients with congenital anomalies. For 9 classes the NHDS ALOS was shorter by amounts ranging from 0.4 days to 15.1 days—the latter for patients with certain causes of perinatal morbidity and mortality. These two extremes were the only classes for which the differences exceeded 1 day.

Most discrepancies in NHDS four-digit diagnostic information arose when there was agreement on principal diagnosis, but disagreement on the appropriate code (referred to as a coding discrepancy). About half the coding discrepancies occurred

when the diagnosis on the NHDS abstract was taken directly from the face sheet of the medical record. In these instances the NHDS coders may have correctly coded the face-sheet diagnosis. However, the face-sheet information was incomplete when compared with more refined diagnostic information contained elsewhere in the record. Although disagreement in the selection of the correct principal diagnosis was somewhat less prevalent (referred to as an ordering discrepancy), it again reflected face-sheet deficiencies and usually was caused by designating the first-listed diagnosis on the face sheet as principal. The relative importance of ordering and coding discrepancies was reversed with three-digit comparisons. However, inadequacies of the face sheet remained the most important contributing factor. Inadequacies in the face sheet led to similar problems in the other studies as well.

Analysis of Principal Procedures

The IOM and NHDS abstracts agreed on principal procedure in about 71 per cent of the cases—a slightly lower figure than in previous studies (Table 6). No procedure was coded for more than two thirds of the original abstracts in any study. For those cases, the reliability was higher than for those cases for which a procedure had been coded.

TABLE 6. Comparisons Between the IOM Field Team and Original Data Source on Principal Procedure (Coded to Three Digits)

	Weighted Per Cent With No Discrepancy		
	NHDS Data	Medicare Record	Private Abstract
All cases	71.4	78.9	73.2
Procedure coded	60.1	56.6	66.0
No procedure coded	76.3	89.7	86.7

The reliability of individual NHDS categories of procedures is shown in Table 7. These figures range from a low of 31.9 per cent for obstetric procedures to a high of 98.0 per cent for proctologic surgery.

Most discrepancies on principal procedure arose from disagreement over whether a procedure was important enough to warrant abstracting and coding. This, in turn, was caused by the field team's strict adherence to the UHDDS definition, which caused them to include more procedures than appeared on the NHDS abstract. (The frequency of this type of discrepancy should diminish as the revised UHDDS definitions and guidelines are implemented.[11]) Procedures most likely to be coded by the IOM and omitted by NHDS were those involving contrast radiography, radiotherapies and "other nonsurgical procedures." The other major reason for discrepancy was the failure of the NHDS abstracter to go beyond the face sheet, which frequently did not contain a procedure, to identify a principal procedure found elsewhere in the medical record.

Adequacy of the Face Sheet

In addition to considering the face sheet as a factor contributing to discrepancies, the field team explicitly considered the adequacy of the face sheet for abstracting every information item, regardless of whether a discrepancy occurred. This information is shown in Table 8. The face sheet usually was adequate for nonmedical data, with the exception of patient disposition. The face sheet was much less adequate for medical information. Face sheet information on principal diagnosis was adequate for 47.3 per cent of the records reviewed; the comparable figure for information on principal procedure was only 5.9 per cent.

The location within the medical record of the correct information on patient dis-

TABLE 7. Discrepancy Between the NHDS Abstract and IOM Reabstract and the Correct Data Source for Categories of Procedures Coded to Three Digits by NHDS (Weighted Per Cent)

NHDS Procedures	No Discrepancy	Correct Data Source If Discrepant				
		NHDS Abstract	IOM Re-Abstract	Either	Neither	Total
Neurosurgery	59.9	—	37.2	2.9	—	100.0
Ophthalmologic surgery	64.0	6.5	16.4	13.1	—	100.0
Otorhinolaryngologic surgery	73.5	2.3	14.1	10.1	—	100.0
Thyroid/parathyroid/thymus or adrenal surgery	75.8	7.2	17.0	—	—	100.0
Vascular/cardiac surgery	56.4	2.2	34.9	5.3	1.2	100.0
Thoracic surgery	69.6	—	30.4	—	—	100.0
Abdominal surgery	82.6	1.8	11.4	4.2	—	100.0
Proctologic surgery	98.0	—	2.0	—	—	100.0
Urologic surgery	70.9	—	21.8	7.3	—	100.0
Breast surgery	62.3	—	24.7	13.0	—	100.0
Gynecologic surgery	58.8	1.1	26.8	13.0	0.3	100.0
Obstetric procedures	31.9	16.8	24.9	26.4	—	100.0
Orthopedic surgery	51.1	1.7	35.2	10.9	1.1	100.0
Plastic surgery	54.8	—	23.2	22.0	—	100.0
Oral/maxillofacial surgery	67.4	14.1	14.6	3.9	—	100.0
Dental surgery	59.3	—	25.2	15.5	—	100.0
Biopsy	54.5	—	35.5	10.0	—	100.0
Diagnostic endoscopy	41.2	—	47.3	11.5	—	100.0
Diagnostic radiography	57.9	13.7	25.9	2.5	—	100.0
Nonsurgical procedures	72.9	14.4	10.6	2.1	—	100.0

position, as well as the medical items, is shown in Table 9. For almost half the cases where the face sheet was inadequate for abstracting patient disposition, the correct information was found on the discharge summary; physician notes were next in importance. The discharge summary contained correct information on principal diagnosis for slightly more than 60 per cent of the comparable cases. The operative report contained the correct information on principal procedure for 52 per cent of the cases where the face sheet had been inadequate; it was followed in importance by

TABLE 8. Adequacy of the Face Sheet for Abstracting NHDS Data (Weighted Per Cent)

Data Item	Adequate	Insufficient	Inaccurate	Total
Admission date	98.8	0.8	0.4	100.0
discharge date	99.2	0.6	0.2	100.0
Date of Birth or Age	96.9	0.8	2.3	100.0
Sex	97.4	1.0	1.6	100.0
Race	91.7	1.0	7.3	100.0
Marital status	95.8	0.2	4.0	100.0
Principal expected source of payment	97.7	0.4	1.9	100.0
Additional expected source of payment	95.9	0.1	4.0	100.0
Patient disposition	24.8	12.0	63.2	100.0
Principal diagnosis	47.3	49.3	3.4	100.0
Principal procedure	5.9	90.6	3.5	100.0

TABLE 9. Location of Correct NHDS Information When Face Sheet Was Inadequate (Weighted Per Cent)

Portions of Record Used	Data Item		
	Patient Disposition	Principal Diagnosis	Principal Procedure
Admit notes	0.1	2.2	—
Consultations	—	0.8	1.1
Discharge summary	48.3	61.4	18.7
Emergency room reports	—	—	0.3
History and physical	0.3	9.8	1.4
Nursing notes	19.3	0.3	1.3
Physician notes	30.9	6.2	7.7
Operative report	—	10.4	52.0
Pathology report	—	1.9	0.3
Other reports	1.1	7.0	17.2
Total per cents	100.0	100.0	100.0

the discharge summary and "other" reports.

Discussion

A review of the findings from all three IOM studies makes obvious the fact that the 1977 data have not significantly improved beyond those examined from 1974. These studies were performed during the time period when four-digit coding of diagnoses was customary. However, the more recent move to a five-digit coding scheme with the introduction of ICD-9-CM, with as yet unestablished levels of reliability, leads to even greater concerns about the reliability of routinely gathered hospital discharge information.

The study committee noted that some improvements in the reliability of hospital discharge data will require fundamental changes in the way that physicians designate diagnoses, the methods by which diagnostic information is classified and the priorities assigned to developing and maintaining good medical record systems within hospitals. These are relatively long-term improvements. However, other improvements can be achieved immediately with relatively simple changes in the processes by which information is abstracted, the design of abstracting forms and their instructions.

Although the current face sheets of medical records are adequate for abstracting most nonmedical information, their inadequacies for abstracting medical data and information on the disposition of the patient at the time of hospital discharge have been clearly established. Significant improvements could be attained if the discharge summary were routinely used to abstract information on patient disposition and principal diagnosis and if the operative report were routinely used to abstract information on principal procedure. Beyond this immediate step, conceptual and methodological improvements are required to further refine all three data items.

With the exception of deaths, information on patient disposition is generally missing from the face sheet and must be inferred from the discharge summary and other portions of the medical record. This information should become a regular component of the face sheet. The concept of patient disposition should be reconsidered and better guidelines developed for its use. If it is intended to describe a patient's continuing health care needs and anticipated service utilization, the current definition

will be misleading in several situations. For example, a patient may be discharged to home for a few days because of the nonavailability of a skilled nursing bed, to which he will be admitted as soon as possible. The current discharge categories would result in a discharge to home rather than, more appropriately, to a skilled nursing facility. Similar ambiguity may exist between patients who are discharged home with no continuing care needs and those who will be served by a home care program.

In the absence of significant improvements in the face sheet and discharge summary, it is unlikely that the quality of codes on diagnoses and procedures will attain optimal improvement unless reference is made to the body of the medical record at the time that codes are assigned. Referring to a narrative summary only is inadequate. Accordingly, the committee recommended that NCHS and other data gatherers, who currently code centrally, experiment with requesting coded information directly from hospitals or purchasing coded abstracts from abstract services known to have data of acceptable quality.

If the split between abstracting and coding functions must be maintained, quality-control activities should refer back to the original medical record to determine the extent to which coded information accurately reflects the condition of the patient. Merely checking the internal consistency of the abstracting and coding as separate functions is not adequate.

Reliability of hospital discharge data is associated with the general organization, orderliness and logic of the medical record. The development of a standardized format and guidelines for compiling the face sheet would be one step toward improving the quality and comparability of medical records.

Unreliable hospital discharge data begins with the assignment of a low priority to medical records systems. Hospitals should increase the resources devoted to medical records activities to produce better basic data, particularly in view of the increasingly important decisions being based on such data. Formal communication links should be developed between persons abstracting data and those dealing with the implications of decisions based on hospital discharge data.

Problems of validity or consistency of diagnostic information recorded by the physician suggest that the undergraduate medical curriculum should place greater emphasis on the probabilistic nature of clinical diagnosis and decision making, the extent to which diagnostic styles vary among physicians even apart from issues of diagnostic acumen and accuracy, current classification schemes and the importance of record keeping.

Part of the cause of unreliable diagnostic data is ambiguity in the criteria for designating diagnoses. Improved diagnostic criteria should be developed, preferably through the consensus development activities sponsored by the National Institutes of Health.

Research should be initiated to develop more valid diagnostic classification schemes to acknowledge and appropriately reflect the presence of co-morbid conditions.

Research should begin immediately to assess the reliability of five-digit coding. Any modifications to ICD-9-CM developed for NCHS, Medicare, or any other routine reporting should be reviewed and approved by the U.S. Committee on Vital and Health Statistics to encourage uniform coding and comparability among all hospital discharge abstract systems, both public and private.

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Hospital Medical Staff Organization and Quality of Care:

Results for Myocardial Infarction and Appendectomy

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This article examines the relationships among hospital structural characteristics, individual physician characteristics, medical staff organization characteristics and quality of care for two conditions: acute myocardial infarction and appendicitis. Using data obtained from the Commission on Professional and Hospital Activities (CPHA), approximately 50,000 acute myocardial infarction cases and 8,183 appendectomy cases collected from 96 hospitals in the East North Central Region of the country (Illinois, Indiana, Michigan, Ohio and Wisconsin) were examined. These data were merged with medical staff organization and related data on hospital characteristics obtained from the American Hospital Association. The results indicate that such medical staff organization factors as involvement of the medical staff president with the hospital governing board, overall physician participation in hospital decision making, frequency of medical staff committee meetings and percentage of active staff physicians on contract are positively associated with higher quality-of-care outcomes, independent of the effects of hospital and physician characteristics. Further, the medical staff organization factors appear to be somewhat more strongly associated with higher quality-of-care outcomes than the hospital and physician characteristics. For acute myocardial infarction, higher volume of patients treated per family practitioner and internist and presence of a coronary care unit were also associated with better outcomes. Given the restricted number of conditions studied, the geographically limited sample and the fact that specific variables were not consistently related to quality of care for both conditions, the results are viewed as preliminary. However, they are consistent with and extend other developing findings in this area. They also suggest that more attention needs to be given to the organization of the hospital medical staff and its articulation with the overall hospital decision-making structure and process in attempts to improve outcomes of hospitalization.

THE RELATIONSHIP of hospital medical staff characteristics and organization to costs and quality of care is receiving increasing attention.[1–6] Factors found to be associated with quality of hospital care include hospital structural characteristics such as bed size and teaching status; individual physician characteristics such as specialty and board-certification status; and medical staff organization characteristics such as degree of staff participation in hospital decision making and coordination and control exerted through committees. Unfortunately, most studies have failed to

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consider all three sets of factors (hospital structure, physician characteristics, medical staff organization) simultaneously in order to examine the relative influence of each.[7] The issue is important because while overall hospital structure and individual physician attributes are relatively difficult to change, medical staff organization processes are much more susceptible to interventions which might improve the quality and efficiency of hospital care.

The purpose of the present study is to examine simultaneously the relationships of hospital, physician and medical staff variables for two conditions: acute myocardial infarction (AMI) and appendicitis. Of particular interest is the examination of whether the medical staff organization variables exert an independent association net of the hospital and physician variables. The existence of such an independent association would suggest that medical staff organization plays an important role as the interface between the hospital and the physician in relation to the quality of care provided. The lack of such an independent association would suggest that the hospital and physician variables are adequately reflecting medical staff organization factors, at least in relation to the quality of care provided for the two conditions under study. If this were the case, it would suggest that little attention needs to be given to medical staff organization itself and that more attention ought to be given to the types of physicians granted staff privileges and the structural characteristics of hospitals granting such privileges.

The above central relationships are examined, taking into account other factors which might be associated with differences in the quality of care provided for the two conditions under study. These include the volume of procedures performed, the degree to which physicians concentrate their work at one hospital, the degree of hospital commitment to quality assurance activities and, in the case of AMIs, the impact of treating patients in a coronary care unit.

The study is primarily exploratory but is guided by a number of propositions suggested in the literature. Given the current state of knowledge, no differential predictions are made for AMI versus appendicitis. Specifically, it is hypothesized that bed size, number of interns and residents per bed and presence of a director of medical education will each be positively associated with higher quality of care. This is based on the assumption that, other things being equal, larger hospitals and those involved in teaching activities have more resources to provide higher quality of care. Board certification as a proxy measure for physician competence is also expected to be associated with higher quality of care. For AMIs it is expected that a higher ratio of internists to family practitioners will be positively associated with higher quality of care. This is based on the assumption that, on the average, internists receive more extensive training in cardiology than family practitioners. Similarly, for appendectomies it is expected that a higher ratio of surgeons to family practitioners will be associated with higher quality of care, assuming that surgeons have more extensive training in surgery than family practitioners. However, in both the AMI and appendectomy cases, physicians' *experience over time with the condition* may offset differences in training.

Based primarily on the findings of Luft et al.,[8] it is also expected that greater volume of procedures performed will be associated with higher quality of care. Given conflicting evidence in the literature,[9–12] no specific prediction is made regarding the effect of treating patients in a coronary care unit, but the available data will enable this relationship to be examined while controlling for a greater number of competing variables and explanations than is presently the case.

The percentage of active staff physicians on contract is hypothesized to be positively associated with higher quality of care, based on the assumption that such a measure at least partially reflects a greater struc-

turing of medical staff organization. Roemer and Friedman[4] have previously found such structuring to be associated with higher quality of care. Concentration of professional activity at one hospital is also felt to be positively associated with higher quality of care on the assumption that such a situation provides more opportunity for peer review, sharing of information among physicians and systematic quality assurance activities. Flood and Scott,[1] for example, have found that the greater the percentage of a surgeon's practice that is concentrated in a given hospital, the better the quality of care.

Physician participation in hospitalwide decision making, measured by such variables as having the medical staff president vote on the governing board and physician participation in the governing board executive committee activities, has also been suggested to be positively associated with higher quality of care.[6,13] The assumption is that such involvement increases the opportunity for awareness and feedback regarding physician practices and provides the opportunity for physicians, administrators and board members to discuss patient care quality issues. Similarly, greater communication via committee meetings among medical staff members is expected to be associated with higher performance. This is because such meetings provide opportunity for learning and peer reinforcement for taking corrective action. Finally, the degree of a hospital's commitment to quality assurance activities is also expected to be positively associated with higher quality of care. Whether or not the hospitals subscribed to an additional medical audit program of the Commission on Professional and Hospital Activities (CPHA) was used as a proxy behavioral indicator of such additional commitment.

Data and Methods

Sample Selection

A 35 per cent proportionately stratified random sample of all hospitals in the East North Central Region (Illinois, Indiana, Michigan, Ohio and Wisconsin) who subscribe to the Commission on Professional and Hospital Activities (CPHA) was selected for study. The East North Central Region was selected because it represents the area of the country within which the highest percentage of hospitals (approximately 45 per cent) subscribe to CPHA. The sample was stratified by bed size, ownership and teaching status, using the following categories: <100 beds, 100–399 beds, and ≥ 400 beds; short-term voluntary general versus governmentally owned hospitals; teaching hospitals (defined as being affiliated with a medical school or having a residency or internship program) versus nonteaching hospitals. These categories were used in combination to create 12 potential strata. In fact, no hospitals existed in the stratum of less than 100 beds, short-term voluntary general and teaching, nor in the stratum of ≥ 400 beds, governmental and nonteaching. Of the 131 hospitals sampled, 96 (73.3 per cent) agreed to release their CPHA data for analysis. There was a somewhat higher nonresponse rate for hospitals under 100 beds but little difference by teaching status or ownership.

Data

Given the problems involved in an attempt to measure the overall quality of hospital care, particularly in regard to adjusting for differences in case mix severity, we chose a *condition specific* approach, examining coronary conditions involving myocardial infarction, ischemic heart disease or disorders of heart rhythm (H-ICDA 410.0–416.9) and conditions involving primary appendectomy (H-ICDA 47.0). Of 54,904 coronary care condition patients, 3,745 were excluded due to missing or misrecorded information, leaving 51,159 records available for analysis.

In order to correct for the fact that some patients who were discharged in a few days may have been miscoded as having an

AMI, we excluded those patients discharged alive with approval in five or less days. We note that currently some patients with documented AMI might be discharged as early as the fifth day, but very few such patients would have been discharged this early in 1973, the relevant year of study. These exclusions represented three per cent of all AMI patients.

The remaining group is referred to as the "filtered AMI patients" and represents 97 per cent of the total AMI patients in the study. In order to adjust the expected mortality for each hospital according to case severity on admission, these patients were stratified based on admission systolic blood pressure level and age, two variables which have strong correlation with mortality and which were available from the CPHA data set.[14–16] The six categories based on age and systolic blood pressure readings upon admission to the hospital were as follows: 1) <65 years old and <90 mmHg; 2) ≥65 and <90 mmHg; 3) <65 and 90–99 mmHg; 4) ≥65 and <90- 99 mmHg; 5) <65 and ≥100 mmHg; and 6) ≥65 and ≥100 mmHg. The number of patients and the number of deaths within each of the six groups were calculated for each hospital, and a standardized mortality ratio (SMR) was developed, indicating the number of expected versus actual deaths for each hospital based on each hospital's number of patients falling into the respective age and systolic blood pressure categories (see technical appendix for calculation).

For the primary appendectomy study, information on 8,695 patients was originally collected for analysis, but 512 patients were excluded due to missing or misrecorded information, leaving 8,183 patients available for analysis. Since we were interested in the pathologic state of the appendix at operation, we stratified patients based on age and sex, as these characteristics are known to be correlated with the probability of the appendix being normal or perforated.[17] Accordingly, the patients were stratified into the following four groups: 1) males ≤5 years old and ≥65; 2) males, 6–64; 3) females ≤5 years old and ≥65; and 4) females 6–64 years of age. For each hospital, information was available on the total number of patients, number of deaths, number of patients with diagnosis of acute appendicitis with peritonitis (H-ICDA code 540.1) and number of patients with tissue result of no abnormality. We chose to focus on the proportion of normals removed rather than of perforations because removal of a normal is an act that can be avoided without necessarily increasing perforation rates. Furthermore, there are major pitfalls in using case rates of perforation, rather than population-at-risk rates, as an indicator of quality.[18] Thus, a standardized normal tissue removed (SNTR) ratio was computed for each hospital in a fashion similar to that of the standardized mortality ratio for AMI patients. (See appendix.) Thus, the quality-of-care indicators for both the AMI patients and the appendectomy patients have been adjusted for differences in case mix severity and standardized against the experience of other hospitals in the sample.

The hospital, physician and medical staff data are derived from a 1973 national survey of hospital medical staff organization conducted by the American Hospital Association. This survey gathered a variety of data on speciality composition, board-certification status, contractual relationships between physicians and hospitals, governance, decision making, committee structure and related items. For 1973, this information was linked with the CPHA data described previously for the sample of East North Central hospitals to conduct the analyses described below.

Methods

The relative effects of the hospital, physician and medical staff characteristics on the AMI standardized mortality ratio and the standardized per cent normal tissue removed were assessed using

multiple-regression analysis. In addition to assessing the effects of the individual variables, indexes of "resource capability," "participation in decision making" and "local staff orientation" were constructed based on previous research and factor analysis of the present data.[6] The variables composing each factor are presented in the appendix. In brief, the resource capability factor represents primarily the ability of the hospital to attract qualified staff and is heavily influenced by hospital bed size and teaching capability. Consistent with previous predictions, it was hypothesized that this factor would be positively associated with better outcomes. Participation in decision making includes measures of physician involvement in hospital wide decision-making bodies and is predicted to be positively associated with better outcomes. Local staff orientation primarily reflects smaller hospitals with fewer physicians per bed and fewer physician members per committee and is hypothesized to be associated with poorer quality of care.

Results are presented for 95 of the 96 hospitals, since data on the SMR were not available for one hospital. In addition, for approximately 12 hospitals, data on the percentage of active staff board certified were estimated using the overall sample mean for per cent of physicians board certified. This substitution yielded results virtually identical to those obtained from deleting these hospitals from analysis. Thus, the substitution results in no bias for the sample overall, and the findings presented below are based on all 95 hospitals.

Results

Acute Myocardial Infarction Results

Table 1 shows the AMI mortality ratios within each of the six age/systolic blood pressure categories. The data indicate significant independent effects of age and systolic blood pressure as well as a strong joint effect. Briefly, controlling for systolic blood pressure on admission, those 65 and over are two to three times more likely to experience death from AMI. Controlling for age, those with systolic blood pressure ≤89 mmHg are approximately three to four times more likely to experience death from AMI than are patients with systolic blood pressure ≥100 mmHg. Looking at the age and systolic blood pressure in combination, patients 65 and over with systolic blood pressure ≤89 mmHg are more than eight times as likely to experience death than those under 65 with systolic blood pressure ≥100 mmHg.

TABLE 1. Mortality Ratios for Filtered AMI Patients Based on Age and Systolic Blood Pressure for 96 Hospitals in East North Central Region

Systolic Blood Pressure	Age	
	< 65 Years	≥ 65 Years
≤ 89 mmHg	0.470	0.828
90–99 mmHg	0.282	0.641
≥ 100 mmHg	0.096	0.334

Table 2 presents the descriptive statistics for all variables examined; the results of the AMI regressions are presented in Table 3. Most of the associations among the variables were in the range of 0.10 to 0.30, with the exception of the correlation between hospital bed size and per cent of active staff physicians board certified ($r = 0.41$) and the association between bed size and existence of a director of medical education ($r = 0.53$). As shown, the volume of AMI patients per internist and family practitioner, having the president of the medical staff on the hospital governing board and hospital size are most strongly associated with lower AMI standardized mortality ratios. The unstandardized regression coefficients indicate that a one-unit increase in volume of AMI patients per physician is associated with a 0.014 decrease in the SMR, controlling for or holding constant the other variables in the analysis.

TABLE 2. Descriptive Statistics for Study Variables

Variable	Mean or %	Standard Deviation	N
Standardized AMI mortality ratio	1.06	0.31	95
Standardized per cent normal tissue (appendectomy) removed ratio	1.00	0.68	96
Hospital bed size	242	171.3	96
Number of interns and residents per bed	2.1/100	5.2/100	96
Director of medical education	25.8%	44%	96
Ratio of internists to family practitioners	1.05:1	2.00:1	96
Ratio of surgeons to family practitioners	1.6:1	2.2:1	96
Number of AMI patients per internist and family practitioner*	5.43	4.51	96
Number of appendectomies per surgeon and family practitioner	2.83	2.10	92
Per cent active staff board certified	43.5%	24.4%	84
Per cent active staff on contract	16.8%	11.8%	96
Per cent staff with major professional activity at this hospital	85.4%	24.0%	96
Physicians on executive committee of the hospital governing board	22.1%	41.7%	96
President of Medical Staff on hospital governing board	38.5%	48.9%	96
Number of meetings per medical staff committee per year	11.64	3.52	96
Resource Capability fac....†	0.082	3.39	96
Participation in decision making factor†	0.001	2.08	96
Local staff orientation factor†	0.091	4.45	96
Hospital reports having a coronary care unit	63.2%	48.5%	96
Hospital also subscribes to CPHA medical audit program	76.0%	42.9%	96

* This is computed by dividing the total number of AMI patients by the total number of internists and family practitioners on the active staff. Ideally data specific to each internist and family practitioner treating AMI patients are desired but were not available from the existing data. The same is true for number of appendectomies per surgeon and per family practitioner.

† See appendix for further description.

Further analysis also revealed that hospitals seeing 60 or fewer filtered AMI patients per year had a standardized mortality ratio of 1.17, as compared with 1.00 for those seeing more than 60 patients per year ($F = 7.06$, $p \leq 0.009$).

The results also indicate that hospitals having their medical staff president on the governing board experienced standardized mortality ratios 0.108 lower than those of hospitals whose staff president was not on the board, again controlling for other variables in the analysis. Presence of a coronary care unit and per cent of physicians board certified were also associated with a lower AMI standardized mortality ratio, but these relationships were not statistically significant. The per cent of physicians concentrating their activity at one hospital was not found to be significantly associated with lower AMI standardized mortality radios. Whether or not the hospital subscribed to CPHA's additional medical audit program and the ratio of internists to family practitioners were not associated with the AMI standardized mortality ratio. Overall, the above variables explained 17 per cent of the variability in the AMI standardized mortality ratios ($F = 3.10$; $p \leq 0.008$). The amount of variation explained by each variable is largely unaffected by the order in which it entered the equation. The exception is board certification, which when entered last in the equa-

TABLE 3. Individual Variable-Regression Results for AMI Standardized Mortality Ratios

Independent Variable	Unstandardized Coefficient (Standard Error)	Standardized Coefficient*
Volume of acute MI patients per family practitioner and internist	−0.0136† (0.007)	−0.200
President of Medical Staff is on Hospital Governing Board	−0.108‡ (0.063)	−0.171
Hospital bed size	−0.00031‡ (0.0002)	−0.170
Physicians concentrate most of their professional activity at one hospital	0.0017 (0.0014)	0.123
Hospital has a coronary care unit	−0.092 (0.071)	−0.144
Per cent board-certified physicians	−0.0016 (0.0016)	−0.121
$R^2 = 0.174$	$F = 3.10, p \leq 0.008$	
$\overline{R}^2 = 0.12$ (Adjusted R^2)		

* The standardized regression coefficients place all variables on equivalent units of measurement based on the standard deviations involved. The magnitude of these coefficients is one way of reflecting the relative importance of the variables in the analysis.

† $p \leq 0.05$.

‡ $p \leq 0.10$.

tion explained only an additional 0.008 per cent but which in the stepwise regression explained 8 per cent of the variation. The magnitude of the standardized regression coefficients indicates the relative importance of the variables, suggesting that volume, followed by president of staff on governing board and bed size, are the most important predictors.

The impact of the overall factors, shown in Table 4, reveals a moderately strong effect of physician participation in hospital decision making on AMI standardized mortality ratios. The unstandardized regression coefficient indicates that a one-unit increase in this factor is associated with a 0.033 decrease in the AMI standardized mortality ratio. Similarly, a one-unit increase in the volume of AMI patients per internist and family practitioner is associated with a 0.013 decrease in the standardized mortality ratio, and the presence of a coronary care unit is associated with a 0.116 lower AMI standardized mortality ratio. The local staff orientation factor was associated with a higher SMR as predicted, but was not statistically significant. Resource capability was not related to the AMI standardized mortality ratio, nor did it make any difference whether or not the hospital subscribed to CPHA's additional medical audit program. None of the other variables of interest were significantly related. Overall, 18 per cent of the variability in AMI standardized mortality ratio was explained by the above variables ($F = 4.17$; $p \leq 0.004$).

When the participation in decision-making factor was forced into the equation first, it explained 7 per cent of the variation in standardized mortality ratios. When it was forced into the equation last, after allowing the other variables to explain as much of the variation as possible, the participation in decision-making factors still explained an additional 5 per cent of the

TABLE 4. Composite Factor Regression Results for AMI Standardized Mortality Ratios

Independent Variable	Unstandardized Coefficient (Standard Error)	Standardized Coefficient
Physician participation in hospital decision making factor	−0.033* (0.015)	−0.225
Volume of acute MI patients per family practitioner and internist	−0.013† (0.007)	−0.187
Hospital has a Coronary Care Unit	−0.116‡ (0.077)	−0.183
Local staff orientation factor	0.010 (0.0085)	0.147
$R^2 = 0.176$	$F = 4.17, p \leq 0.004$	
$\overline{R}^2 = 0.134$		

* $p \leq 0.025$.
† $p \leq 0.05$.
‡ $p \leq 0.10$.

variation. An interaction term involving volume times presence or absence of a coronary care unit was also examined but was not significantly related to the standardized mortality ratio.

Appendectomy Results

Table 5 shows the per cent of normal tissue removed within each of the age/sex categories. As indicated, age makes no difference for males, but females between the ages of 6 and 64 are more than twice as likely to have normal tissue removed as females 5 years or younger and 65 or older. The per cent of females between the ages of 6 and 64 with normal tissue removed is approximately twice that of males 6 to 64. The higher rate for females 6 to 64 is primarily due to the number of female conditions, such as ovarian cyst and pelvic inflammatory disease, whose symptoms may be closely related to those of acute appendicitis.

TABLE 5. Per Cent Normal Tissue Removed by Age and Sex for 96 Hospitals in East North Central Region

Age	Male	Female
≤ 5 and ≥ 65	0.088	0.067
6–64	0.083	0.145*

* The increased rate for females is primarily due to the number of female conditions, such as ovarian cyst or pelvic inflammatory disease, whose symptoms may be closely related to those of acute appendicitis.

The results shown in Table 6 are based on 92 hospitals, due to missing data on the variable volume of appendectomies per surgeon and family practitioner. The findings also exclude an "outlier" hospital,* which performed nearly twice as many appendectomies per surgeon and family practitioner per year than other hospitals in the study (21.67 vs. 12.83) and had a standardized per cent normal tissue removed ratio of 3.13, as compared with 2.61 for the next-highest hospital. As shown, the frequency of medical staff committee meetings, the degree to which physicians concentrated most of their activity at the hospital under study, the per cent of physicians on contract and the presence of a director of medical education are each significantly

* A 322-bed voluntary community general hospital with no interns and residents.

TABLE 6. Regression Results for Standardized Per Cent Normal Tissue Removed

Independent Variable	Unstandardized Coefficient (Standard Error)	Standardized Coefficient
Frequency of medical staff committee meetings	−0.057* (0.021)	−0.280
Physicians concentrate most of their professional activity at one hospital	−0.0068* (0.003)	−0.240
Percentage of active staff physicians on contract	−0.013† (0.006)	−0.222
Hospital has a director of medical education	−0.246‡ (0.162)	−0.155
Volume of appendectomies per family practitioner and surgeon	0.028 (0.037)	0.088
$R^2 = 0.17$	$F = 3.56$; $p \leq 0.006$	
$\bar{R}^2 = 0.12$		

* $p \leq 0.005$.
† $p \leq 0.025$.
‡ $p \leq 0.10$.

associated with a lower standardized per cent of normal tissue removed. The unstandardized regression coefficients indicate that a one-unit increase in the frequency of staff committee meetings is associated with a 0.057 decrease in the SMR; a one-unit increase in per cent of physicians concentrating their activity at one hospital is associated with a 0.0068 decrease in the SMR; a one-unit increase in the per cent of active staff physicians on contract is associated with a 0.013 decrease in the SMR; and the presence of a director of medical education is associated with a 0.246 decrease in the SMR. Whether or not the hospital subscribed to CHPA's additional medical audit program, the ratio of surgeons to family practitioners, number of appendectomies performed per surgeon and family practitioner, per cent of physicians board certified and other variables included in the analysis were not associated with the standardized per cent normal tissue removed ratio. The lack of a volume/quality relationship for appendectomies confirms the results of Luft et al.[8] Overall, frequency, of committee meetings, concentration of professional activity at one hospital, per cent of physicians on contract and presence of a director of medical education explained 17 per cent of the variability in the per cent standardized normal tissue removed ratio ($F = 3.56$; $p \leq 0.006$).

It should be noted that the variables were not sensitive to the order in which they entered the equation. For example, whether forced in first or last, volume explained none of the variation in per cent standardized normal tissue removed. In contrast, frequency of committee meetings explained 7 per cent of the variance even when forced in last and also explained 7 per cent when forced into the equation first.

Analysis of the overall factors involving resource capability, medical staff participation in decision making and local staff orientation revealed no significant associations. The only variable that continued to be significantly associated with a lower standardized per cent normal tissue removed ratio was concentration of major professional activity at the hospital under study. The higher the percentage of physicians indicating that they concentrated

their activity at one hospital, the lower the per cent standardized normal tissue removed ratio.

Discussion

Differences in mortality in coronary care units are well documented.[19-21] Less well documented are the factors that might be associated with these differences. Bloom and Peterson found that coronary care units with full-time medical directors experienced better performance.[19] Stross et al.[21] found that units seeing fewer than 60 AMIs a year experienced *twice* the mortality of those seeing 60 or more per year. Others have identified the critical elements in treating CCU patients, including the positive impact of educating CCU nurses in problem recognition and response.[22,23] Policy changes permitting nurses to intervene immediately have also been associated with declines in mortality.[24] On a different level, the very effectiveness of CCUs in treating AMI patients has been questioned relative to at-home treatment and treatment in regular hospital beds.[9-12]

The present results suggest that the presence of a CCU does make a difference in the standardized mortality ratios of AMI patients. Hospitals with a CCU had lower standardized mortality ratios than those without a CCU, even after controlling for hospital size, volume and related variables. However, for those hospitals with a CCU, it was not possible to analyze whether those patients in the CCU did better than those treated in private or semiprivate rooms.

Further, the results suggest that volume of AMI patients treated is, indeed, a significant factor, with those treating 60 or fewer AMI patients a year experiencing a 17 per cent greater standardized mortality ratio than those treating more than 60 a year. More relevant is the fact that number of AMI patients treated *per* internist and family practitioner is significantly related to lower standardized mortality ratios, even after taking other variables into account. This finding suggests that recent research linking volume of surgical operations to postoperative mortality may also be part of a more generalized phenomenon in which the amount of clinical experience with a given condition within a given year plays an important role in nonsurgical cases as well.[8] If supported by other studies, the results suggest some important implications for grouping or clustering CCU services around those hospitals with sufficient volume to maintain high performance outcomes. Under the guidance of health systems agencies and faced with other external pressures, hospitals are already beginning to group or cluster their obstetric and pediatric services. However, treatment of AMI patients raises a somewhat more difficult set of issues in terms of travel time to receive treatment factors. Thus, a policy of reducing the number of individual hospitals with CCUs would also depend in part on the availability of competent emergency medical transport services available, particularly in rural and semirural areas. The question also arises as to whether volume or the existence of a CCU is the more important factor. Results from Table 4 suggest that both are about equally important in terms of their association with standardized mortality ratios. However, analysis of covariance involving only volume, presence or absence of a CCU and standardized mortality ratios revealed a stronger effect of volume ($F = 7.91$; $p \leq 0.006$) than of presence or absence of a CCU. ($F = 2.70$; $p \leq 0.104$.)

The present results involving AMI standardized mortality ratios also suggest the important role played by medical staff organization, particularly in regard to participation in hospital decision making. Whether or not the president of the medical staff was on the hospital governing board and an overall measure of participation, which included also whether or not

other medical staff officers were on the governing board, and whether there were physicians on the governing board executive committee, were significantly related to lower standardized mortality ratios, taking into account other possible explanatory factors. In fact, the present results suggest that the overall medical staff participation in decision making factor is the single most important variable associated with lower standardized mortality ratios, more significant than either the volume of AMI patients per internist and family practitioner or the presence or absence of a coronary care unit. Other variables commonly believed to be associated with better performance, such as hospital bed size, degree of teaching activity and per cent board-certified physicians, were not as significantly associated with lower standardized mortality ratios when the above variables were taken into account. Each of them, however, was in the expected direction.

The literature suggests wide variation in the removal of histologically normal appendices.[25-28] Yet, again, there is little documentation of the factors that might be associated with the amount of normal tissue removed. Neutra's research suggests that with a more careful history and exam and close observation for several hours, hospitals can reduce the rate of normal tissue removed without any increase in the rate of perforation.[25] The present analysis suggests some medical staff organization factors which may be associated with and may facilitate the kinds of behavior suggested by Neutra. Specifically, frequency of medical staff committee meetings, concentration of physician activity around one hospital, per cent of active staff physicians on contract, and presence of a director of medical education are each associated with a lower standardized per cent normal tissue removed ratio. Frequent medical staff committee meetings promote communication and exchange of information among staff members, increasing the probability of implementing changes in treatment practices, which may improve the process and/or outcomes of care. Lending support to this notion, Morlock et al. recently found a significant association between frequency of medical staff committee meetings and lower overall hospital death rates after adjusting for patient diagnostic mix.[29]

Concentration of professional activity at one hospital offers greater opportunity for peer review and establishment of new practice habits and patterns. For example, in early quality-of-care research, Morehead found that those physicians who spent most of their time (greater than 75 per cent) treating Health Insurance Plan (HIP) of Greater New York patients provided a higher quality of care than those scattered throughout several locations or those seeing HIP patients only part-time.[30] Flood and Scott have found that the greater the percentage of a surgeon's practice that is concentrated in a given hospital, the better the quality of patient care.[1]

Having a greater percentage of physicians on contract also affords an opportunity for closer physician involvement in hospital treatment practices and facilitates the opportunity to initiate and implement improvements in care. The present results in this regard are also supported by previous work which found positive associations between measures of physicians on contract and a variety of quality-of-care measures.[4] Finally, the presence of a director of medical education may provide the kind of systematic leadership and direction which, when combined with the greater emphasis on teaching implied by such a position, may stimulate better treatment practices.

The above findings and interpretations are only suggestive. Variables associated with differences in outcomes for AMI were not the same as those associated with differences in outcomes for appendicitis, although both sets of variables contained

measures of medical staff organization concepts. Findings for these conditions should be replicated. Subsequent research should also examine other conditions as well as hospitalwide measures of quality in order to extend present results. Further specification of volume–quality relationships with regard to concentration of admissions and operations around smaller numbers of physicians would be a useful avenue of inquiry. Particularly useful would be studies of the impact of planned changes in medical staff organization and hospital–physician relationships on quality of care that would include data from hospitals not introducing such changes. Variations in implementation of quality assurance programs might be a source for such studies. The impact of regulation (rate review, certificate of need, and so on) on physician–hospital relationships and medical staff organization in relation to both cost and quality considerations is also a fertile area for study. Finally, it is of interest to note that the relatively aggregate measures used in the present research explained approximately 15 per cent of the variation in quality of care. Thus, examination of medical staff and hospital organization at a less aggregate or more micro level (such as the actual organization and staffing of coronary care units) would be likely to add to the explanation of differences in quality of care across hospitals.

While inferences drawn from the above results can, in a strict sense, only be generalized to other CPHA-subscribing hospitals in the East North Central Region of the country, the findings are suggestive of some important medical staff influences which may hold implications for many U.S. hospitals. Of particular interest is the fact that, for the conditions and hospitals studied, neither hospital structural characteristics such as bed size and degree of teaching involvement nor individual physician characteristics such as specialty composition were strongly related to better performance. Rather, medical staff organization characteristics, such as degree of physician participation in hospital decisionmaking, frequency of committee meetings, concentration of activity in one hospital and percentage of physicians on contract, were found to be more strongly associated with superior performance, although not consistently so for both conditions. In addition, for AMI patients, the presence of a coronary care unit and the volume of AMI patients per internist and family practitioner were related to superior performance. It appears that simply having "good" people (e.g., well-trained, board-certified physicians) and having these people affiliate with "good" hospitals (e.g., large-size, teaching-oriented hospitals) is not sufficient to guarantee superior results. Rather, the actual organization of the medical staff itself, as reflected partly by participation in hospitalwide decision-making bodies and communication through committee meetings, appears to be most strongly associated with the quality of care provided. Thus, activities aimed at changing medical staff organization and articulating its relationship and involvement with the overall hospital organization and decision-making process may be a key factor in improving the quality of hospital care.

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Appendix

A. Calculation of Standard Mortality Ratio (SMR) for AMI Patients and Standardized Normal Tissue Removed Ratio (SNTR) for Appendectomy Patients.

1. Standardized Mortality Ratio (SMR) =

$$\frac{\sum_{i=1}^{6} d_i}{\sum_{i=1}^{6} n_i \ (P_i)}$$

where:

d_i = each hospital's number of AMI deaths in each age, systolic blood pressure category

n_i = each hospital's number of filtered AMI patients in each age, systolic blood pressure category

P_i = total number of deaths in each age, systolic blood pressure category for *all* study hospitals *divided by* the total number of patients in each age, systolic blood pressure category for *all* study hospitals.

2. Standardized Normal Tissue Removed Ratio (SNTR) =

$$\frac{\sum_{i=1}^{4} t_i}{\sum_{i=1}^{4} n_i \ (P_i)}$$

where:

t_i = each hospital's number of patients with normal tissue removed in each age, sex category

n_i = each hospital's number of appendectomy patients in each age, sex category

P_i = total number of patients with normal tissue removed in each age, sex category for *all* study hospitals *divided by* total number of appendectomy patients in each age, sex category for *all* study hospitals.

3. *Factor Analysis Results*

1. *Resource Capability Factor*

Variables	Factor Loadings	Communality
Hospital bed size	0.82	0.82
Existence of a director of medical education	0.67	0.55
Number of services and benefits provided to medical staff	0.51	0.51
Per cent of active staff board certified	0.45	0.63
Number of interns and residents/bed	0.36	0.77
Per cent of active staff who are radiologists and pathologists	−0.41	0.81

Eigenvalue = 4.96.
Per cent of variance explained = 30.3 per cent.

2. *Participation in Decision Making Factor*

Variables	Factor Loadings	Communality
Other medical staff officers are on governing board	0.84	0.81
Physicians are on executive committee of governing board	0.44	0.37
President of medical staff is on hospital governing board	0.29	0.41

Eigenvalue = 1.39.
Per cent of variance explained = 8.5 per cent.

3. *Local Staff Orientation Factor*

Variables	Factor Loadings	Communality
Per cent of active staff who are radiologists and pathologists	0.82	0.81
Per cent of active staff who are generalists (i.e., family practitioners and pediatricians)	0.32	0.29
Number of active staff physicians per bed	−0.31	0.90
Per cent of active staff board certified	−0.36	0.63
Number of medical staff members per committee	−0.36	0.64
President of medical staff is elected by the staff	−0.43	0.30

Eigenvalue = 1.60.
Per cent of variance explained = 9.8 per cent.

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Communications

Noninvestigational Use of Unapproved Drugs—Experience From the Scandinavian Countries

ELINA HEMMINKI, M.D.

In the Scandinavian countries it is possible to use unapproved drugs for treatment purposes with special permission from the drug control authorities (individual licenses). The Scandinavian practice and the problems connected with it are described. The basic principle underlying the laws and regulations has been and is the same in each country: unapproved drugs should be used only in exceptional cases for special reasons. However, the practice has become different: many individual licenses have been granted, giving physicians much greater opportunity to prescribe drugs of their choice than the apparently tight Scandinavian licensing laws and practices would suggest. The great number of individual licenses has become an administrative problem and has also undermined the principle underlying the law.

THE PRESENT DRUG LAWS in the United States allow interstate sales only of those drugs which have been approved (licensed) by the Food and Drug Administration (FDA). However, according to the report of the U.S. Senate Committee on Labor and Human Resources,[1] FDA, under the guise of approving drugs for investigational purposes, has permitted small numbers of patients with serious diseases to use unapproved drugs. The 1979 Drug Regulatory Reform Act[1] proposes that with certain criteria, use of unapproved drugs should be allowed in the treatment of small numbers of patients.

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In the Scandinavian countries it has traditionally been possible to use unapproved drugs for treatment purposes with special permission from the drug control authorities. (These permissions will be called "individual licenses" henceforth.) The purpose of this article is to describe the Scandinavian practice and the problems connected with it.

The data on which the article is based consist of laws, statutes and other official documents; archive material; interviews and previous studies. (A list of laws and other official documents is available from the author.)

The Laws and Practice

Laws requiring drug licensing came into force in 1930 in Norway, in 1934 in Sweden, in 1954 in Denmark and in 1964 in Finland. The general features of the present licensing laws and practices have been

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described earlier.[2,3] In Denmark and Sweden the option for individual licenses was stated in the laws from the beginning of drug licensing. In Norway this option was introduced in 1941, and in Finland in 1969. But at least in Finland, individual licenses were permitted before that date.

The basic contents of the laws and regulations have been and are the same in each country: unapproved drugs should be given only for special reasons, and to obtain such drugs a special application must be sent to drug control authorities. In Finland, Norway and Sweden the special reasons must be health or health care related. In Norway and in most years in Sweden (1934–1952, 1963–present), it has been further stated that the reasons should be given for each patient. In the other years in Sweden, individual licenses were permitted for individual drugs for specified physicians, who could use them for several patients. In Finland and Denmark the laws have been more vaguely formulated, and whether individual licenses should be given per patient, physician or institution has not been clearly specified.

The laws and other official documents give an impression of strict rules, but the practice has been different. In each country applications for individual licenses have been accepted more or less automatically, and only a few have been rejected. One or two people have handled the applications for individual licenses, usually in addition to their other duties. Comparison with the large number of applications shows that consideration of each application cannot have been thorough. In Norway until 1977, there was not much difference for physicians between writing an application for an individual license and writing a normal prescription. Only since 1977 has a special statement justifying the

TABLE 1. Number of Individual Licenses in the Scandinavian Countries*

Year	Denmark	Finland	Norway†	Sweden‡
1949	NA	NA	?	2,100
51	NA	NA	?	3,242
55	?	NA	?	3,086
57	?	NA	2,120	7,539
59	?	NA	2,433	9,798
61	?	NA	3,280	20,486
63	?	NA	3,200	23,525
65	?	?	3,272	11,013
67	?	?	3,984	3,133
69	?	?	4,284	3,417
71	?	?	4,738	6,141
73	?	?	6,546	14,922
74	960	917	7,757	19,502
75	973	872	~ 9,800	20,803
76	1,668	1,244	~12,400	23,350
77	1,971	1,424	16,189	27,802
78	?	1,959	17,618	33,409
79	?	?	12,929	39,594
Population in 1978, in 1000s	5,104	4,752	4,059	8,278

* The Danish and Finnish figures concern number of applications for individual licenses, the Norwegian and Swedish figures licenses granted. NA = not applicable; ? = not available.
† Includes individual licenses for drugs used in clinical trials.
‡ Includes individual licenses for special foods, see text.
Sources: 4–12.

need of the unapproved drug been required. Pharmacists have taken care of the application for the individual license: they have ordered the drug, and a copy of the prescription has been sent to the drug control authorities, who have decided about an individual license. In the other countries the practice has been similar, but statements about need were already required as long ago as 1956 in Sweden.

The Danish, Finnish and Norwegian laws have stated that only small amounts of drugs should be given through each individual license. In Finland this was explicitly stated by the drug control authorities: for outpatients it was limited to the amount necessary for 3 months of treatment between 1969 and 1975, and to a year since 1975, while amounts for hospital use could be larger. Nowadays, most of the individual licenses in Finland are given to hospitals. The allowable amounts have not been defined as exactly in the other countries. Yet, apparently, at least those individual licenses given to hospitals and not specified per patient have permitted large amounts of drugs. Such unspecified individual licenses for hospitals have also been given in Norway and Sweden, even though they are against the spirit of the regulations about individual licenses. In Norway such licenses have accounted for about a quarter of all individual licenses in the 1970s.[4] It is true that in Norway the drug control authorities have been given a general title to exempt the licensing regulations. Apparently, this title has been applied when these unspecified individual licenses have been granted.

The numbers of individual licenses show that they have not been rare exceptions (Table 1). The figures are those given by the control authorities themselves. They are not comparable between the countries, because the types of products covered by individual licenses and the amounts allowed on one license vary from one country to another. For example, many individual licenses for special foods are included in the Swedish data, and herbals and homeopathics are included in the Norwegian data. In Sweden there were 19 individual licenses for special foods (0.6 per cent of the total) in 1967, 2,816 (19 per cent) in 1973, and 12,848 (46 per cent) in 1977.[13] On the basis of the data given by Christiansen,[4] I estimated that in 1977 licenses for herbals and homeopathics accounted for a fifth of all individual licenses in Norway. In the statistics one individual license may refer to a small amount of a drug for a short treatment of one patient or to a large amount for the treatment of several patients in a hospital. (The individual licenses given to specified physicians in the period of 1952–1963 in Sweden were apparently recorded per patient, and this explains the high number of individual licenses during that period.) Within each country the recording practices do not vary much in a short time period, and it may be concluded from Table 1 that in each country the number of individual licenses has increased during the 1970s. In 1979 the control authorities in Norway had to handle as a mean 60 requests per working day, and in Sweden some 180 requests.

Reasons for Individual Licenses

I have interpreted the spirit of the Scandinavian drug laws and determined that the following have been "acceptable reasons" for individual licenses: 1) treatment of a disease which occurs only rarely in the country; 2) continuation of treatment of a person only visiting the country; 3) use of a drug regarded as undesirable for general use but considered beneficial for some patients. As "unacceptable reasons" the following three can be listed: 4) an application for the drug to be licensed has been made, and individual licenses are applied for to pressure the control authorities and to promote the drug; 5) the control authorities have rejected or would reject an application for the drug, so individual licenses are used to avoid conflict (a means of getting around the licensing regulations); 6) the producer of the drug does not

want to market the drug for commercial or technical reasons.

The numbers of individual licenses and especially their increase suggest that the acceptable reasons cannot explain all of them. Part of the increase in the individual licenses in Sweden may be explained by the licenses for special foods (see above). But even when these are excluded, the number of individual licenses has been high and is increasing in Sweden. A study by Christiansen[4] allowed me to estimate the reasons for individual licenses in Norway in the 1970s. She took a one-third random sample of individual licenses in 1974 and 1977. In her report the drugs for which individual licenses were most commonly applied (66 per cent of all applications in 1974 and 79 per cent in 1977) are listed and discussed. I reanalyzed her raw data and grouped the individual licenses by the status of application for a general license (Table 2). Table 2 can be interpreted and simplified in the following way: about one third of the individual licenses may have represented pressure on control authorities and premarketing promotion (group 1). Another third may have been for the purpose of getting around licensing regulations (group 2). (The individual licenses for herbals and homeopathics are included in this group, because I believe that the Norwegian control authorities have considered them ineffective.) These are crude estimates, and a small proportion of the individual licenses in groups 1 and 2 have been for acceptable reasons. About 15 per cent of the individual licenses were for drugs which the producers did not want to market (group 3). The actual numbers of individual licenses increased from 1974 to 1977 in each group, because the total number more than doubled during this period.

Discussion

The large number of individual licenses, the limited resources reserved for their handling within control authorities, the individual licenses to hospitals which were not per patient, scattered data from several sources and my analysis from the Norwegian data (Table 2) show that the individual licenses in the Scandinavian countries have not reflected their original purposes. They have given physicians a much greater

TABLE 2. Individual Licenses by the Status of the Application for General License; Most Commonly Applied Drugs in Norway in 1974 and 1977*

	%			
	1974		1977	
1. Application under consideration	32		35	
2. Application rejected	12	34	12	37
Drug withdrawn from market by control authorities' decision	1		2	
No application; drug likely to be ineffective†	21		23	
3. Drug withdrawn from market by decision of the firm	2	14	1	16
No application, drug likely to be effective	12		15	
4. No application, efficacy unknown	13	20	9	12
Cannot be classified	7		3	
Total	100		100	
Number of individual licenses	(1605)		(4175)	

* Based on the data provided by Christiansen (1979), see text. Drugs prescribed by dentists and veterinarians are excluded.

† Includes herbals and homeopathics.

opportunity to prescribe drugs they choose than the apparently tight Scandinavian licensing laws and practices would suggest. To the drug industry it has offered a way of putting pressure on the control authorities while the application for licensing is under consideration and a way of promoting the drug before approval. I do not know whether firms have received direct financial profits from individual licenses. In Norway individual licenses accounted for 2.4 per cent of the value of the total sales of drugs in 1977.[5,14]

In recent years the administrative burden caused by the increasingly large number of applications for individual licenses has drawn attention to them. To alleviate the situation a proposal has been made in Sweden to allow certain institutions to use unapproved drugs officially without applying for permission for each patient (a practice widely used in the other Scandinavian countries).[13] In Norway new administrative measures and information aimed at reducing the number of individual licenses have been put forward, and since 1977 the number of individual licenses has started to decrease.

Yet the great number of individual licenses is of concern also as a matter of principle. The purpose of drug licensing in the Scandinavian countries is to ensure that only effective, safe and reasonably priced drugs are on the market. Certain bodies within the drug control authorities have been charged with making these judgments. The system of individual licenses was created to allow exceptional use of unapproved drugs for special (health-related) reasons. If these principles are considered poor they should be officially changed, and not breached in practice so that part of the drug consumption is determined by a different principle, i.e., the judgment of practicing physicians.

My conclusion is that the principle of the individual licenses could be retained, but that present practice should be changed to correspond to the spirit of the law, with a reduction in the numbers of individual licenses. A special problem is posed by individual licenses for useful drugs, which drug firms for one reason or another have not tried to bring onto the market. Soliciting of applications for such drugs, or for unpatented drugs, their production by noncommercial institutions, might help.

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